Real House Price Dynamics in OECD countries

The risk of large movements in prices

Mari O. Mamre

Thesis for the degree
Master in Economic Theory and Econometrics

Department of Economics

UNIVERSITY OF OSLO

February 2014

Real House Price Dynamics in OECD countries

The risk of large movements in prices

2014

Real house price dynamics in OECD countries –the risk of large movements in prices

Mari Olsen Mamre

http://www.duo.uio.no/

Trykk: Reprosentralen, Universitetet i Oslo

Preface

It must have been about 9 months ago I decided to ask one of my favorite professors in macroeconomics, Asbjørn Rødseth, about his ideas for a master thesis project. Discussion went back and forth, as it has ever since. I do not remember exactly whose idea it was to focus on house prices. What I do remember, is the personal engagement, knowledgeable inputs and useful supervision he has put into this project. For that I am ever grateful. It goes without saying this thesis would not have been the same without him.

I also wish to thank the Centre of Equality, Social Organization and Performance (ESOP) at the University of Oslo for financing further work based on this thesis in collaboration with Asbjørn Rødseth. The frame of a master thesis is a bit too narrow to go all the way in addressing some questions raised by this analysis.

An acknowledgement must also be made to Ragnar Nymoen who always took the time to answer every small and not so small question I had about methodological issues, be it in the corridor passing through, after lectures or me running down the door at his office.

Finally I am grateful for all the inspiring discussions with other students at the Department of Economics. A special thanks to Even Winje for his assistance with software-related issues.

Any errors or inadequacies are on the writers behalf.

Table of contents

List of tables

List of figures

1 Introduction and summary

Developments in real house prices differed substantially across OECD countries over the period 1975-2012. In the majority of OECD countries, real house prices have been moving up over the last decades. As can be seen in Figure 1 below where we have plotted the real house prices for 11 OECD countries, some countries, such as Ireland, Luxembourg and Australia, experienced protracted price appreciations while other countries prices were more stable over time, such as Germany. Most countries prices display a cyclical pattern, some more than others, as the cases of Spain and Italy. The weighted aggregate for 21 OECD countries portraits an upward trend until the financial crisis started about 2007, with one additional apparent break in this trend around year 1991 (bottom right panel in figure 1).

The consequences of the collapses of the housing markets for the real economy in many countries in both the 1990's and 2007/2008 cases have been severe. When homeownership-rates are high and households are highly leveraged, increases in systemic risk can bear consequences for large parts of the population. There are non-trivial differences in the homeownership-rates across OECD countries. At the high end of the scale are Spain and the UK (83.2 and 70.7 percent were homeowners in 2004, respectively (OECD numbers)), while at the low end were Germany and Switzerland (only 41 and 38.4 percent owned a home). One would thus expect busts to affect a broader part of the population in the high end countries, leading to adverse effects among different segments of the population.

Using different econometric approaches and based on a panel of 21 OECD countries this thesis investigate whether differences in structural or policy factors significantly affects the price responsiveness of shocks to demand in the short run and in the cases of abrupt movements in real prices. Over such steeper areas of the housing cycle the analysis focus specifically on finding evidence of asymmetric responses of demand and structural factors on price dynamics. The study of asymmetries in the overall business cycle is a well-developed field in econometric analysis but, to my knowledge, no such efforts are put into the study of housing cycles. This thesis will not go into the concept of over-evaluation or misalignments in the housing market, which are related concepts.

Catte et al. (2004) investigate the relationship between house prices and the overall business cycle for OECD countries between 1975 and 2004. They find that property markets tend to track the business cycle, with a tendency for real house price turning points to lag business cycle peaks. This pro-cyclical nature of housing prices has led others to study the feedback effects to the macro economy from both consumption and wealth effects. When consumers have an infinite planning horizon, sharp increases in prices today may lead to an increase in consumers' current valuation of permanent income (Friedman 1957), leading to increased borrowing today in order to smooth consumption over the life-cycle. Moreover, housing wealth can facilitate access to credit for liquidity-constrained households (Catte et al. 2004 p.14). Higher prices also stimulate investment in new dwellings, depending on the elasticity of supply in the respective country.

This thesis is inspired by new insights on the connections between shocks to demand, structural and policy factors and the effect on house price dynamics and the work done by Andrews (2010) in particular, as well as the idea of asymmetries over the business cycle. In certain respects the analysis borrows heavily from choices of structural factors and shocks to deem relevant for housing demand from Andrews. Also, the aim is to bring something new to the table. Whereas Andrews consider long term determinants of prices and overall volatility, I use a similar framework to consider short term dynamics of prices as well as medium term periods when prices have risen or fallen abruptly (booms and busts) separately. I have also built a richer dataset for 21 OECD countries between 1975 and 2012. Some alternative econometric approaches will also be discussed.

I find evidence of different influential driving factors involved for a simple predictor of large house price falls and large house price increases. The existence of asymmetries in the driving factors behind housing cycles could make demand relations dependent on a sort of switching regime, in which case any prediction problem become a nonlinear one. This would evidently affect empirical results based on symmetric models. In a symmetric short term price dynamics model these results indicate the income elasticity of housing prices decrease if the degree of banking supervision is high. However, evaluating periods of large movements in prices, this effect dissipates in periods of booms and increases the risk of a bust. Moreover, according to these estimates, rigidities in supply is an important driving factor in a booming period but not influential during periods of a bust.

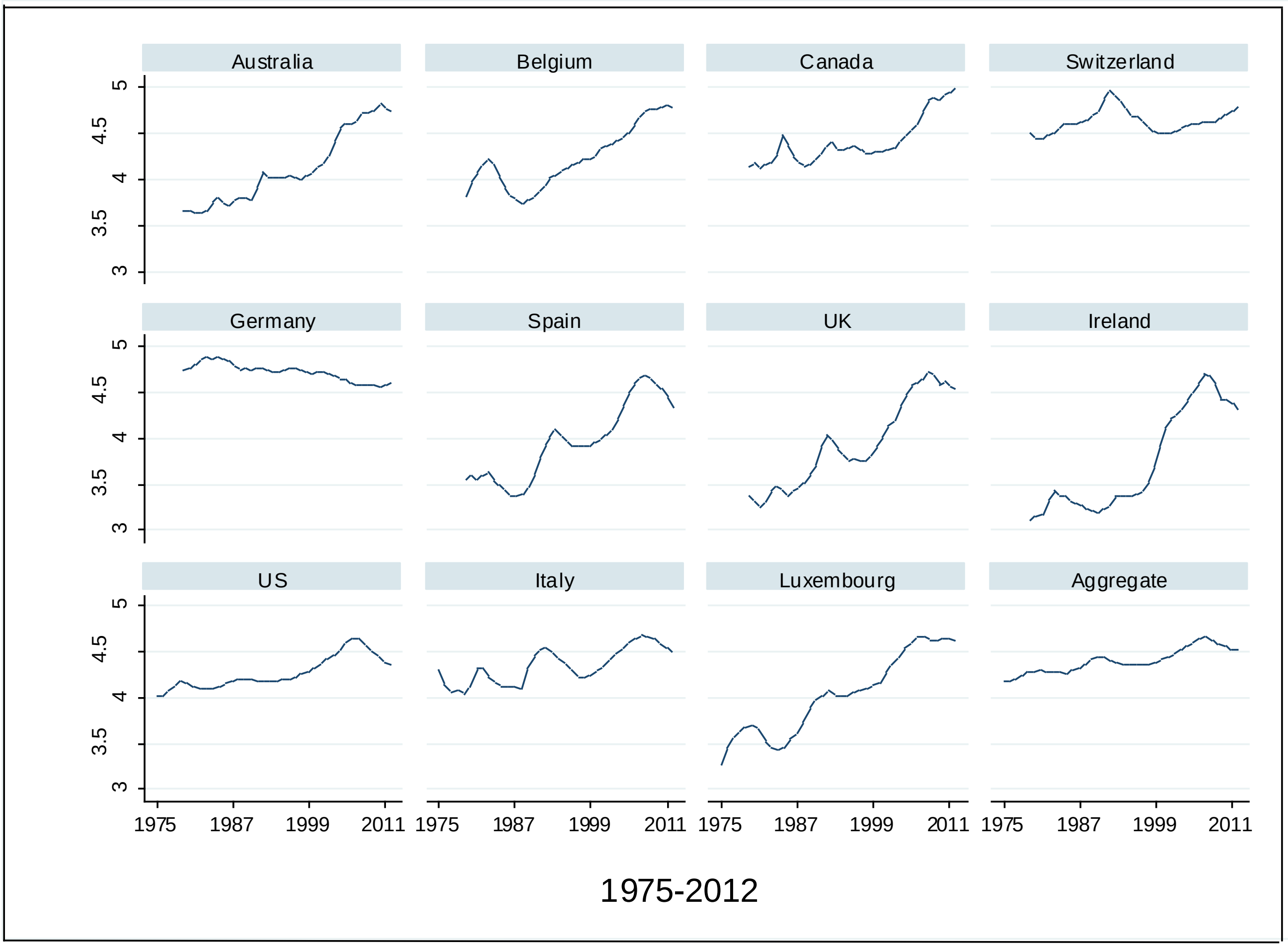
Considering contraction periods only, an in-sample forecasting model of price falls in a similar framework as the more contemporaneous models discussed in the previous section is estimated. Evaluating the models' predictive abilities at historical episodes of large price falls, the risk probabilities are in all but a couple of cases much higher than the country specific average at the eve of a bust. For Belgium, Canada, France, the Netherlands ('79 case) and the US ('79 case) the calculated probabilities are between 2.5-3.2 standard deviations away from their respective means.

Authors at the Bank of International Settlements (BIS) have recently (2013) compiled a monthly database for policy actions on mortgage and housing markets for numerous countries between 1990 and 2012. With the luxury of hindsight, the overall impression after reading through the various policy actions is that many measures were taken subsequent to a recent downfall in the housing market or when the private sector were especially highly leveraged. This could suggest a "leaning against the wind"-attitude towards build-ups in both markets. Of course, policy makers often face a trade-off between acting towards such build-ups and at times conflicting goals such as battling unemployment or securing the competitiveness of the exporting sector.

The linkage of the housing market to the real economy, the business cycle and the potentially adverse consequences of a bust calls for attention towards the determinants of house price dynamics, both in the short, the medium and in the long run. There has been a newborn focus on structural and policy factors' effects on real housing prices among especially OECD economists (see *inter alia* Andrews 2010 and OECD Economic Policy Reforms 2011). If there is an accelerator mechanism present in housing markets (Bernanke 1996), knowledge of which policy measures or structural reform dampens an upswing and mitigates a downturn is warranted. And if there are asymmetric driving factors over the housing cycle, this would call for differences in responses.

The organization of the thesis is as follows; Section 2 discusses theory of price formation and develops a model for the housing market. Section 3 describes the data. Section 4 lays out the econometric approach and estimation results for short term price dynamics with some extensions. Section 5 consider large price movements and proposes some simple cut-off values for booms and busts and estimation results with a particular focus on abrupt price falls. At last we summarize some main findings in section 6.

Figure 1.1 Log real house prices by country 1975-2012. Index: 2005=100



Data: Federal Reserve Bank of Dallas, US.

2 Theory of the housing market

This section presents a model of the housing market using the concept of general equilibrium that will form the theoretical basis for the analysis in the rest of this thesis. I then turn to review some theory of determinants of housing demand and supply and their possible linkage with structural country factors. It will also be clarified what variables that is used in the econometric analysis and their interpretation.

2.1 A model of the housing market

The theoretical equilibrium framework adopted is inspired by the stock-flow model of the housing market as described by DiPasquale and Wheaton (1992) and Sanchez and Johansson (2011). Let $\boldsymbol{X} \in \boldsymbol{R}^n$ be a vector of exogenous demand side factors such as income, structural factors, the user cost of holding the housing asset and demographic characteristics and let P the real price of housing. These concepts will be clarified in more detail later. Assume the demand for housing $D(\cdot) : \boldsymbol{R}^n \to \boldsymbol{R}$ determines the equilibrium real price that will clear the finite stock of housing S ∈ ℕ for the owner-occupied housing market in the long run

$$D(\boldsymbol{X}, P) = S := \int_{t=0}^{T} S(t)dt \quad (1)$$

, where $t = 0$ is an arbitrary date in the past. The housing stock is assumed similar in quality. On the supply side, the stock of housing slowly depreciates over time, at a rate $\delta > 0$ and increases with new housing investments *I*, which depend on cost shifters $\mathbf{Y} \in \boldsymbol{R}^n$ such as demographic factors and the costs of production and on real house prices *P:*

$$dS = I(\boldsymbol{Y}, P) - \delta S \quad (2)$$

Assume demand for housing consumption $D(\cdot)$ is approximately continuous and monotone in each of its arguments (a consequence of this assumption is flexible prices which rules out e.g. inertia in prices). Also assume the initial dwelling stock $S(0)$ is predetermined.

Then inverting (1)

$$P_t = D_t^{-1}(\boldsymbol{X}_t;\, S_t) + u_t^D \quad (1)'$$

Any observed period *t* real price can be expressed as a function of demand side factors given the housing stock in place and an error term. If (1) determines the long term equilibrium price the error term in (1)' can be interpreted as shocks that lead the actual price deviate from the equilibrium price

$$P_t^* = D_t^{-1}(\boldsymbol{X}_t;\ S_t)$$

$$P_t - P_t^* = u_t^D$$

When the relationship (1)' is assumed approximately log-linear, where lower-case letters indicates the natural log we get

$$p_t = \alpha_0 + \boldsymbol{\alpha}_1 \boldsymbol{X}_t' + \alpha_2 s_t + u_t^D$$

Considering year on year changes we achieve a short run inverse demand equation of price growth which is our baseline equation of interest

$$\Delta p_t = \beta_0 + \boldsymbol{\beta}_1 \Delta \boldsymbol{X}_t' + \beta_2 \Delta s_t + \varepsilon_t^D \quad (3)$$

Short run since we consider period (yearly) changes. However, in the short run one could argue the dwelling stock be regarded as fixed. By allowing the dwelling stock to vary in (3) and given the proper inclusion of medium term[1] determinants of housing demand in **X,** the relationship can hold for medium term price dynamics as well.

On the supply side one can think of the flow of investments into housing as contributing to both the stock of housing increasing over time and maintaining the existing stock

$$I(\boldsymbol{Y}, P) = dS + \delta S \quad (2)'$$

Assume that investments will depend on period *t* price expectations conditional on the available information set $\boldsymbol{\Omega}$. Then in a similar manner, a long run relationship for housing investment can be expressed as

$$i_t = \gamma_0 + \gamma_1 E_t(p_{t+1}|\boldsymbol{\Omega}) + \gamma_2 \boldsymbol{Y}_t' + v_t$$

For our analysis we are not interested in determinants of supply directly, but as shown by Sanchez and Johansson (2011) the responsiveness of housing supply to price changes is important in determining the evolution of house prices, and particularly the volatility of prices. There are reasons for suspecting interactions between demand and supply.

[1] I follow the conventional understanding of "medium term" as about 2-4 years, depending on the context.

I next turn to review some of the main components or concepts of demand and supply individually in the subsequent paragraphs and the variables used in relation to the theory in the econometric analysis

2.2 Income and credit

Macro-financial linkages 1

A natural starting point in any demand function is income. Price growth in line with growth in households' income can last for a substantial amount of quarters. Both from the perspective of access to borrowing linked to affordability and from the supply side it must be the case that disposable income, within reasonable bounds, constrains house price growth. The relationship between house prices and mortgage markets gained interest in the 1970's when several advanced economies experienced their first house price booms (e.g. UK). Some argued the availability of credit was the most important driving factor, others that income growth was most important.

Meen (2002 p. 13-14) show that credit constraints change the responsiveness of house prices to variables such as interest rates in the UK economy. Furthermore, if also down payment constraints (required deposit to buy a house) exist, then prices will be more sensitive to changes in income than in their absence. Meen use a simple model of this relationship on the UK economy:

$$P = \frac{\gamma W}{(1 - dr)} \qquad dr = \alpha P \tag{2.1}$$

,where P is house price, W is income and dr is the down payment ratio such that $(1 - dr)$ is the 'Loan to Value' ratio available on a mortgage.

The units are chosen such that $dr \in (0,1)$. A short disposition of the arguments is as follows: When $dr = 0$, the elasticity of house prices with respect to income equals 1, as is seen in some empirical studies[2] (long term elasticity). With $dr > 0$, as income rises, house prices rise more than proportionally. The down payment ratio itself rises with the level of prices ($\alpha > 0$). At higher prices, the risk to lenders may rise and they may require a higher down payment. Also, existing owners are wealthier and need to borrow less to finance a new house. Then variations in the down payment *increase the volatility* of prices compared with income changes, for given interest rates.

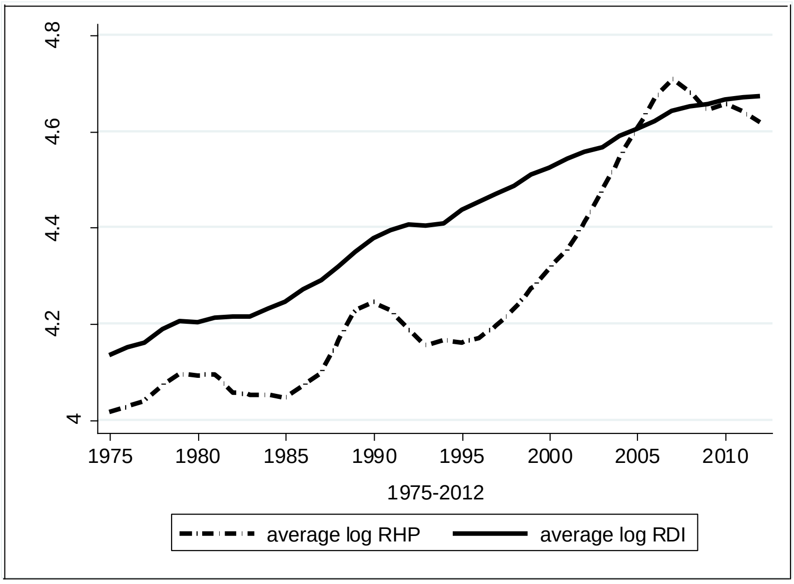
Meen's simple model may not have captured the ensuing development in mortgage markets for many OECD countries. In the US, down payment requirements fell towards zero at the turn of the millennium. In 1989 only 1 in 230 homebuyers made a down payment of 3 percent or less, by 2007 it was 1 in 3 (AEI 2012 p.1) along with substantial increases in house prices when mortgage backed security sales spiked. Within the simple model (2.1), such a feature would indicate that down payment ratios started *decreasing* when prices soared ($\alpha < 0$) from developments outside the model. Although insufficient to describe the interactions between credit and demand, its essence bears some useful insights: Poorly regulated credit markets where a larger proportion of income can be spent on new house buys may amplify prices, which again may increase wealth and then demand (macro-financial linkages).

In the same line of reasoning follow later work by Almeida et al. (2006). Their results indicate that cross-country differences in the maximum Loan to Value (LTV) ratio available on a mortgage can explain cross-country variation in the sensitivity of house prices and credit demand to income shocks. The initial shock changes household wealth through changes in housing prices, this in turn shifts household debt capacity, amplifying the impact of the initial shock on demand and prices, and so forth.

Figure 2.1 plots the evolution of average real house prices (RHP) and disposable income (RDI) for 21 OECD countries.

[2] Terrones and Otrok (2004) estimate an income elasticity of 1.1 over 1970-2003 for 18 countries, OECD Economic Survey (2004) find an elasticity of 1.9 over the same period.

Figure 2.1 Development of log RHP and RDI. Unweighted country mean



Index: 2005=100

Price to income ratios have increased in the average OECD country during the recent run-up and then fallen after a peak around 2007. Price to income ratios are often used as a simple measure of affordability in housing. When prices grow much faster than disposable incomes, it could be signs of a bubble in the housing market. The fact that this has been the case in many OECD countries over 1985-1990 and 1995-2007 is *cet par* compatible with the accelerator theory; yearly increases in income give a proportionally larger increase in prices during the run-up years to the 1991 and 2007 “crisis kick-off”-years

The maximum LTV ratio or down payment can, as discussed, influence the income-sensitivity of prices if there is an accelerator mechanism present. But most countries have a whole battery of regulations toward the mortgage market in tandem with ceilings on maximum loans such as risk-weighing of borrowers and some do not have ceilings at all (BIS housing policy dataset 2013). We want a broader measure of access to credit or bank prudence for cross country comparison. The degree of banking supervision in a country (Abiad et al. 2008) provides such an alternative.

This indicator increases if the country has implemented a capital adequacy ratio based on the Basle standard, if the banking supervisory agency is independent from executives' influence, if a banking supervisory agency conducts effective supervisions and more so if these audits cover all financial institutions without exception. Andrews find that a high degree of banking supervision decreases overall volatility of prices (Andrews 2010 p. 25). We would expect this indicator to affect price dynamics through the income channel from an access-to-credit effect.

2.3 The user cost of housing

Macro-financial linkages 2

A well-developed strand of literature on house price formation is grounded in asset pricing theory where the price of real estate regarded as an asset should arguably be in line with asset returns (see *inter alia* Poterba 1992). Asset returns when considering housing is again closely linked to the costs of holding the asset relative to other investments and expected capital gains (the user cost of housing). Price to rent ratios provide a second measure of disparity and asset price theory suggests that changes in the cost of owning versus renting a house in equilibrium could be used to predict price movements in the owner occupied sector. The user cost (UC) of housing provides this link. The conventional UC as defined in Poterba (1991 p. 1-9) is expressed as

$$UC = \left[(1-\tau_w)\left(i+\ \tau_p\right)+\ \beta+m+d-\pi\right] \qquad (2.2)$$

, where τ_w is the marginal income tax rate, i the nominal mortgage interest, τ_p the property tax rate, $(\beta+m+d)$ constitute *the recurrent holding costs* and consists of the risk premium on residential property (β), maintenance (m) and depreciation (d) and π is the expected capital gain.

For the sake of brevity I will not go into further details of the user cost theory here, but settle with stating that i) the user cost theory of housing and the related concept of permanent income is well founded in microeconomic theory and ii) many influential econometric models of house price dynamics incorporates elements from equation (2.2) (two seminal contributions are Poterba 1991 for a support of the theory and Englund, Ioannides 1992 for a critique).

From the user cost theory this analysis will use the CPI, the general price increase, as a proxy for the expected capital gain, the residential property tax and the real interest rate measuring the opportunity cost of foregone interest or the higher cost of mortgage loans in the econometric models of price dynamics.

The preferential tax treatment of housing may affect price formation through both the supply and demand side. The analysis will consider shocks to structural unemployment being a measure of economic prospects or activity combined with differences in the degree of property tax (macro-financial linkages). Andrews (2010 p. 25-27) find that tax relief on housing increases the overall volatility of house prices and points to the possibility that generous tax policies may increase the expected net profits from speculative investments. It is natural to expect the residential tax to work in the opposite direction as the tax relief.

The most relevant interest rate depends on the mortgage structure of a country. Depending on the composition of countries mortgages between fixed and variable rates I will use real long term and short term interest rates respectively. By increasing the costs of a loan it is expected that higher real rates, all else equal, put downward pressure on prices.

2.4 The formation of price expectations

The user cost theory is however silent as of the *formation* of the expected capital gains from buying a house. This feature may be important for price dynamics. A house is most often household's most valuable asset. When faced with a price increase, given opportunity to buy the house, whether an agent is willing to make the investment could depend on several factors; the duration of the investment and the motivation of the buy, among them. If the house is bought mainly as a consumption good, but the household has a short term perspective before a resale then the expected resale value matters, in line with more speculative investments.

However, there is some evidence that home buyers do not generally seem to be primarily driven by prospects of investment returns. A study by the Building Societies Association (BSA 2007 p.6-12) in the UK indicates that the primary motives for buying a home were the desire to own (88% of respondents) and that rental payments were wasted money (77%).

Financial gains were important for less than half of the respondents (45%). But a concern for would-be homeowners was the risk of being priced out of the market if prices rose (this was a motivation for 28% of first-time buyers and 63% of people considering buying for the first time). Hence, expectations of future price increases *can* increase demand even in the absence of speculators purely motivated by the prospects of capital gains. If expectations are formed on the basis of price history, such dynamics is important. There might of course be country heterogeneity in the formation and importance of price expectations.

2.5 Supply

Supply-demand linkages

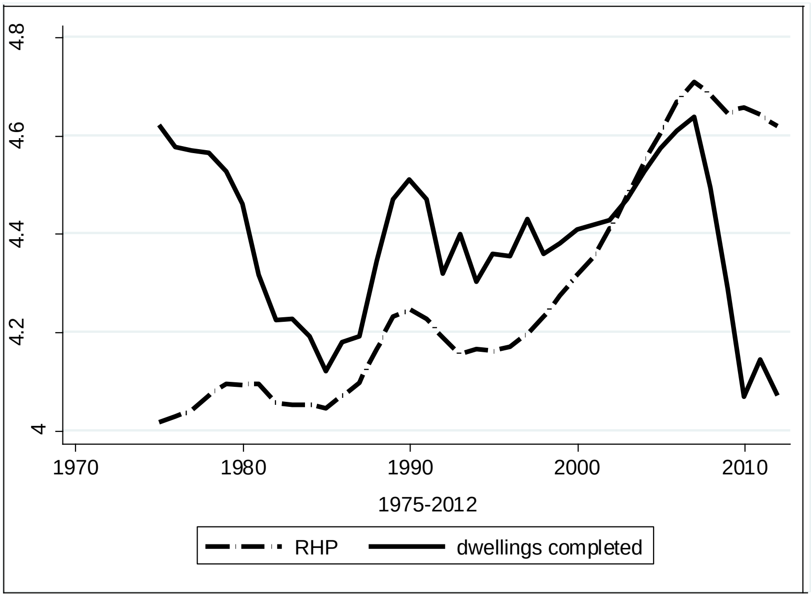
As briefly discussed in 2.1 there may be important supply-demand linkages influencing the responses of any shock to demand on prices. There are large country variations in the responsiveness of housing investment following a price increase. Supply-side restrictions such as zoning laws, the effectiveness of passing through building applications and transaction costs differ among OECD countries.

Sanchez and Johansson (2011) estimate the long run elasticity of investment supply for 21 OECD countries over 1980-2010 in a simultaneous equilibrium model of demand and supply with error correction. According to these estimates, the US and Sweden have the highest elasticity of investments (2.014 and 1.381 respectively) while the Netherlands and Switzerland have the lowest (0.186 and 0.146 respectively). The authors point to the dual effect of flexible supply; more flexible supply may lower the effect on prices when there is a shock to demand but also contribute to more cyclical house price swings in periods of growth through large increases in investment. Since the estimate period and countries are broadly the same, I will employ the results of this study in the subsequent analysis of price dynamics.

Both the price responsiveness of income and the effects of demographic pressure may depend on the elasticity of supply. The dual effect theory then suggests that when demand for housing increases due to either an increase in income or more people competing for houses, higher house supply responses may, everything else equal, dampen the pressure on prices in an upswing since a larger proportion of the adjustments take place through quantity effects. But the mere fact that investments are cyclical may contribute to price swings.

Figure 2.3 plots the evolution of a proxy for the change in the dwelling stock and real prices at the OECD sample mean. Since dwellings completed necessarily lags the decision to invest, the broadly compatible larger turning points of these series suggests that price increases (decreases) lags increases (decreases) in investments, consistent with that investors on average have correct expectations of near future price dynamics which would propose a forward-looking approach.

Figure 2.2 Log dwellings completed and log RHP



Data: OECD and national sources. Unweighted country mean

3 Data description

Using a panel of 18 OECD countries this thesis will investigate ***whether differences in structural or policy factors such as the degree of banking supervision and taxation of property significantly affects the price responsiveness of shocks to demand in the short run (section 4) and in the cases of abrupt movements in prices and if there are evidence of asymmetries (section 5).*** The model is evaluated against episodes of large price falls and lastly a closer look at the run up years before the crash in the US housing market will be taken.

Quarterly data on nominal and real second hand (mainly) house prices and private per working-age population disposable income for 21 OECD countries over the years 1975-2012 are sourced from Federal Reserve Bank of Dallas, Globalization and Monetary Policy Institute and described in Mack and Martinez-Garzia (2011). Accordingly, the authors select closely comparable house price indices which are homogenized as much as possible to make the panel more suitable for comparisons across countries and time. Which market segment that is covered varies between countries[3].

Data on country yearly population density and real interest rates are collected from The World Data Bank (WB). Density is mid-year population divided by square kilometer land area excluding area under inland water, national claims to continental shelf and exclusive zones. This variable is chosen to capture demographic changes and relates demographics to available land (a 1 % increase in highly densed countries may put more pressure on housing availability than low densed ones). Real rates are calculated as $r_t = \frac{(i_t - \pi_t)}{(1+\pi_t)}$, where i_t is period *t* nominal lending interest rate and π_t is period t inflation (GDP deflator).

Annual long term interest rates and the CPI are collected from OECD Economic Outlook No 93 and OECD Monthly Monetary and Financial Statistics. It is 10 years interest rates on government bonds. Real long rates are calculated by subtracting the annual CPI. A composite real rate is constructed using the real long rate for countries where the mortgages are mainly issued with fixed rates and the IMF real rates where mortgages mostly issued with variable

[3] E.g. the Canadian prices cover apartments and two story dwellings located in ten main metropolitan areas whereas German prices cover annual, nationwide existing (second hand) single-family houses (most of the series are more similar than this example). In the econometric analysis, the inclusion of country specific effects will help deal with possible concerns raised by this cross-country heterogeneity in types of real estate.

rates. The information on the composition of mortgages is gathered from The European Mortgage Federation and OECD (Wyman 2003 p.29-30) and OECD Economic Outlook no. 80 (2006 p. 145). Data on the NAIRU (non-accelerating unemployment rate) are sourced from OECD Economic Outlook no. 93 for all countries less S. Africa, which are supplemented by the (ordinary) unemployment rate sourced from IMF World Economic Outlook database. The nairu measures structural (equilibrium) unemployment and are less prone to short term temporary fluctuations not grounded in fundamental sources.

Yearly data for new dwellings started are collected from OECD MEI for 10 countries and the rest of the countries less Luxembourg are supplemented from national sources using the search engine 'Macrobond' and aggregated into years where necessary[4]. This particular variable turned out to be the most problematic one for which to find comparable country data.

Data on private sector credit issued by deposit money banks and other financial institutions to GDP are collected from Beck et al. "Financial development and structure Dataset" (WB, revised 2013).

In the absence of comparable numbers for the overall symmetry or lack of symmetry in taxation of financial objects for our purpose, we use data on the recurrent (annual) taxes on immovable property from OECD tax statistics as percent of GDP. Most countries have an annual property tax through the majority of the years 1975-2012.

Yearly country data on financial reform and the degree of banking supervision between 1975 and 2005 for the countries in our panel (less Luxembourg) are sourced from the financial reform database by Abiad et al. (2008). These are somewhat broad measures of the degree of financial deregulation in a country and the overall index is based on a total of 8 dimensions such as credit controls and reserve requirements, interest rate controls, entry barriers in the banking sector and the degree of banking privatization.

We use the estimates of long term country supply elasticities from Sanchez-Johansson and the indicator measuring the degree of banking supervision in a country from Abiad et al. as discussed in section 2.

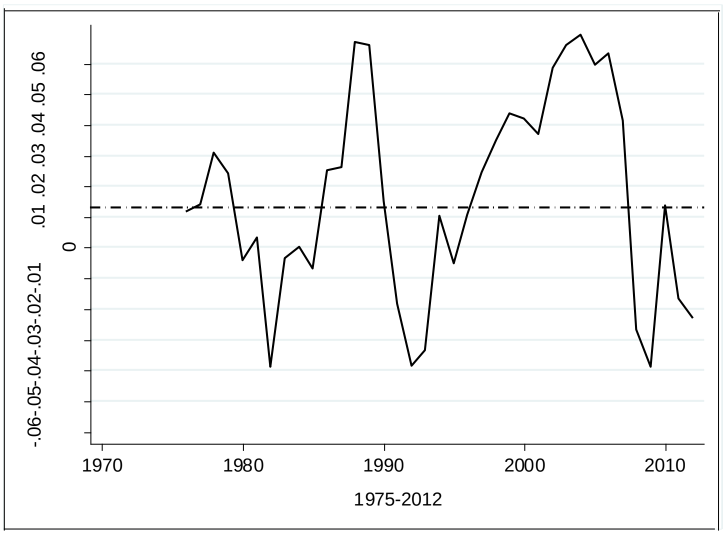
[4] For Italy, Ireland and South Africa we have only data on number of permits issued for residential dwellings, which is lagged one year and used as a proxy. We lack several time-observations for some of the countries.

4 Short term price dynamics - econometric analysis

Throughout I will use the labels “short term price dynamics” or “short term model” for the results and models developed in this section to distinguish these from the models in the next chapter which are labelled “large price movement’s models” or the like in relation to “medium term price dynamics”. This is mainly categorical and do not presuppose that a complete analysis of either short term or medium term house price changes is executed. If so, there would have been a case for e.g. using shorter time periods such as quarters in this section.

To estimate our equation of interest for short term dynamics (3) I use the change in the log real house price index as the dependent variable. There are several arguments for this choice, instead of using levels. Specifically, taking logs will linearize the growth in the original series and stabilize the variance of the differences. Graphically, the sample mean of the differenced log prices[5] seem to revert to a mean about approximately 1-2 % growth with quite persistent deviations, and thus vulnerable to serial correlation[6] (Figure 3.1). Including a constant is then equivalent to assuming a stochastic trend. Note however that this highly aggregated statistic does not give more than a broad picture of the average OECD development.

Figure 3.1 First differenced log RHP.



Data: Federal Reserve Bank of Dallas, US. Unweighted sample mean

[5] Taking second differences did not improve on the first differenced version, by informal comparison of graphs.
[6] First differenced log real disposable income looks about the same evaluated at the sample mean; strongly trended in the logs, more stationary in first differences (Appendix 1)

As a starting point it is useful to check the order of integration of the series House prices and disposable income are found to have strong trend. Our series are quite long (T=38) compared with the number of groups (21 countries), although the panel is of such relative size that the panel aspect is likely to matter. The Maddala and Wu (1999) panel unit root test indicate that log real house prices are $I(1)$. Cook and Vougas (2009) confirm the stationary characteristics of house price changes, but find that prices in addition exhibit structural breaks.

Upon visual inspection, most country series of rates of price gains appear to show evidence of at least two structural breaks; the 90-91 recession looks more common to each country while the financial (housing) crisis 2007/2008 appear more idiosyncratic in nature. To incorporate these features and test for unit roots we use the Clemente-Montañés-Reyes (CMR) unit root test with double mean shifters, additive outliers model[7]. As can be seen in Table A2 in the appendix , in line with previous research mentioned above, we cannot reject the null of a unit root in log prices for all 21 OECD countries in our sample including the weighted composite using a 5 % significance level. Moreover, the two structural breaks are almost everywhere significant, except for 3 cases[8].

Since we are interested in modelling in terms of changes of the variables, rather than levels, there is no need to test for co-integration among the price and the other variables provided the year % changes of income and the dwelling stock (figure 3) are approximately $I(0)$ as well, which for income is confirmed by the Levin-Lin-Chu test (appendix) and for dwellings suffices to examine Figure 2.2 on page 13. I will however, by the way of modelling, investigate the cointegration properties of the model through an error correction model.

There is ample evidence of a consistent autoregressive pattern in data series for housing prices. Several studies find positive autocorrelation at short lags and negative autocorrelations at longer lags, consistent with the boom-bust cyclical patterns (see e.g. Englund and Ioannides 1997 p.11). In this case, also the first differenced variable could be serially correlated in which case inference from estimation could be false. Here it is adjusted for serial correlation using different methods.

[7] The additive outliers (AO) model captures sudden changes in series. Canarella et al. (2010) show including structural breaks in unit root test is superior to other tests using US real estate price indices.
[8] The 3 cases are the estimated breaks for Norwegian prices in 1987, Denmark in 2007 and S. Korea in 2004.

4.1 A baseline model

The aim of this section is to investigate the statistical properties of models, compare these and evaluate estimates of driving factors ***behind short run (yearly) price changes to develop a model of the risk of large movements in prices in the medium term (3 years)***, which is the topic of the next section. There it will be argued that symmetrical models like these may not capture the "true" effects on price dynamics, at least in the steeper areas of the housing cycle. For the moment, however, we direct attention towards developing fruitful models for short term price dynamics and their implications.

The theoretical framework is equilibrium in the housing market where short term price dynamics is explained by the demand side, but where households face different barriers such as more or less prudent banks across countries, as outlined in section 3. The baseline (4.1) models short term house price dynamics in a Time Series Cross Section (TSCS) random effects (RE) framework where changes in the log of real house prices are explained by changes in the log of real disposable income, changes in the dwelling stock, unemployment, the real or long term interest rate and structural factors identified through shocks to demand and a vector of controls for a country *i* at year *t*. The constant term is interpreted as a stochastic trend for the house price level.

$$\triangle\, logP_{i.t} = \beta_0 + \beta_1 \triangle\, logW_{i,t} + \beta_2 \Delta nairu_{i,t-1} + \beta_3 banksup_{i,t-1} \times \triangle\, logW_{i,t} + \beta_4 \Delta logS_{i,t-1} + \boldsymbol{\beta}_5 \boldsymbol{Z}'_{i,t} + \alpha_i + \varepsilon_{i,t} \quad (4.1)$$

$$, t = 1, \ldots, T_i, \quad i = 1, \ldots, N$$

, with AR(1)-errors $\varepsilon_{i,t} = \rho\varepsilon_{i,t-1} + \epsilon_{i,t}$, $|\rho| < 1$, where $\epsilon_{i,t}$~IID(0, σ_ϵ^2) or white noise. Note that it is assumed that the parameter ρ is homogenous across countries.

Alternatively $\varepsilon_{i,t}$ is IID across countries as the two-way (separate year intercepts) fixed effects alternative to (5). We will consider extensions to the baseline model.

Since the panel is long($T > N$), asymptotics are in T with N fixed.

The dependent variable is the yearly percentage change in real house prices[9] for year *t* and a country i, $\triangle\, logW_{i,t}$ is yearly percentage growth in real disposable income (RDI) potentially included with lags and also interacted with structural factors such as the degree of banking supervision in the country which varies a bit but not much over time (but varies between countries) and the price elasticity of supply, $\Delta logS_{i,t-1}$ is the lagged log change in the dwelling stock and $\boldsymbol{Z}_{i,t}$ is a vector of time-varying controls including the start of period price inflation (CPI), % growth in the population density (pop. density) and the (lagged) change in the real rate or the long term interest rate, depending on the composition of mortgages between fixed and variable rates (Composite real rate).

α_i are unobserved country random or fixed effects assumed constant over time and could be country heterogeneity in degree of home-ownership, real-estate structure and other time constant variables[10] not captured by the model. Note that country specific-effects in this model refer to factors affecting the *price changes*. In the FE framework, the unobserved heterogeneity is allowed to be conditionally arbitrarily correlated with the explanatory variables, while the RE framework instead assume that α_i are realizations of an IID process with zero mean and constant variance, that is, the model puts the unobserved heterogeneity in the error term.

I find that the FE model is rejected in favor of the RE approach by the Hausman test for some but not all specifications. Note also that (4.1) can be regarded as an *a posteriori* relationship. Several demand shocks have been tested in relations to structural factors but proved insignificant (more specifically, all the specifications in Table 5.2-3 in section 5 have been tested here as well). Results are reported in section 4.3.

[9]

$$\triangle\, lnP_{i.t} = ln\left(\frac{P_{i.t}}{P_{i.t-1}}\right) \simeq \frac{P_{i.t} - P_{i.t-1}}{P_{i.t-1}} = \%$$

[10] The examples mentioned are not necessarily time constant, but some features of them most probably are. E.g. cultural attitudes toward owning a home and the reported country price series.

4.2 Extensions to the baseline model

One could instead model the dynamics by including a lagged dependent variable (LDV). The bias introduced will be ~ $O\left(\frac{1}{T}\right)$. The contrasts to the AR(1) case will be related to dynamics, as well as interpretation (Beck 2005). The process of adjustment of changes in real prices may depend on the difference between the equilibrium level and previous years actual level or on habit formation, in which case a dynamic model is suited. Theoretical arguments for assuming the change of prices last year explains changes today could be the evidence we discussed briefly in section 2.4 from the study by the Building Societies Association (BSA, 2007) in the United Kingdom where many prospective homeowners report being concerned by the risk of being priced out of the market if prices rose as a rationale for entering the housing market. Then any period *t* demand depend on the previous vector of covariates, the price today and the price last period, which in the long run translates into the equilibrium price

$$D(\boldsymbol{X}_t, P_t, P_{t-1}) = S \quad (1)'$$

In addition we can pursue the equilibrium assumption of house prices further and estimate these models in an error correction framework (ECM). This approach is common in the housing literature (see *inter alia* Annett et al. 2005). If the level version of the demand equation holds in the long run, and prices are originally in equilibrium, any period change in one of the explanatory variables will induce two effects on prices; 1) the short term effect as measured by (4.1) and 2) a deviation from long term equilibrium. In essence it is just another way of modeling dynamics and the error correction theory provides nice interpretation

$$\Delta p_t = \beta_0 + \beta_1 \Delta \boldsymbol{X}_t' + \beta_2 \Delta s_t + \gamma EC + \varepsilon_t^D \quad (4.2)$$

Below are estimation results for the baseline AR(1) model, the model with a lagged dependent variable (LDV) and the error correction version (ECM), all with country random or fixed effects for the shorter time period 1976-2006 for which data on the degree of banking supervision is available. The ECM is estimated by first running OLS on the (time) demeaned level variables using robust errors and then using the lagged residuals (EC) in the differenced version. (L) signify the variable is lagged and panel robust standard errors are in parenthesis. The overall impression when comparing models is that dynamics seem to be important and the FE LDV and ECM models have some advantages over the AR(1)-error model (LR $\chi^2(1) = 18.07$)[11].

It would not, in general, be expected that the structural factors significantly affect short term (yearly) house price dynamics, and the interactions have been included for comparison with the longer term dynamics. However, we get significant effects of the degree of banking supervision combined with income growth (columns 2.1-2 and 4.1-2).

[11] This test statistic compares the AR 1 model to the LDV model. The ECM is just a reparameterization

4.3 Results short term price dynamics models

Lags of income changes are not significant in the baseline (not included in results). All variables are in changes except the structural factors. (L) signify the variable is lagged and standard errors are in parentheses

Table 4.1 Estimation results short term real house price growth AR(1)-model

Dependent variable: RHP yearly % growth. AR(1) Baseline model (1.1-1.2) and shocks to real income (2.1-2.2)

Variables	(1.1) RE_AR1	(1.2) FE_AR1	(2.1) RE_AR1	(2.2) FE_AR1
RDI	0.444***	0.446***	0.852***	0.723**
	(0.0934)	(0.0963)	(0.261)	(0.281)
NAIRU (L)	-0.00540	-0.00418	-0.0549***	-0.0400**
	(0.00391)	(0.00397)	(0.0131)	(0.0158)
Composite real rate (L)	-0.00214**	-0.00247***	-0.00296***	-0.00353***
	(0.000833)	(0.000882)	(0.000881)	(0.000925)
Dwelling stock (L)	-0.0170*	-0.0133	-0.0287**	-0.0124
	(0.0101)	(0.0101)	(0.0112)	(0.0120)
Pop. density (L)	1.272	1.755	2.597***	3.193**
	(0.818)	(1.109)	(0.925)	(1.325)
CPI (L)	-0.00564***	-0.00686***	-0.00374***	-0.00557***
	(0.00115)	(0.00141)	(0.00120)	(0.00149)
Demand shocks and structural factors				
RDI × Banking Supervision			-0.230***	-0.226**
			(0.0857)	(0.0895)
RDI × Supply elasticity			-0.146	0.0113
			(0.238)	(0.258)
Constant	0.0992**	0.0875***	0.151***	0.0858***
	(0.0451)	(0.0171)	(0.0499)	(0.0219)
Total impact (elasticity) of interacted variables on short term price growth at the mean of the structural variables (mean degree of banking supervision= 1.385 over 1975-2006				
1 % RDI growth			0.533	0.410
Observations	493	473	384	366
Number of countries	20	20	18	18
R squared (within)		0.11		0.17
Estimated AR(1) coefficient[12]	0.572***	0.572***	0.582***	0.582***
Mod. Bhargava et al. Durbin-Watson	0.975	0.975	0.979	0.979
Hausman (p-value)	0.0675	0.0675	0.322	0.322

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

[12] Serial correlation is tested for by the TSCS analogue of the standard LM-test. The t-stat of the OLS residuals of the model specification (1) on the lagged residual is 14.95. The residual is not significantly related to the other independent variables, except for the change in real disposable income, making the exogeneity-assumption in this model questionable, at least in the OLS-version. All specifications are tested and *** signify the coefficient on the lagged residual significant at 1 %.

Table 4.2 Estimation results short term real house price growth LDV and ECM models
Dependent variable: RHP yearly % growth. LDV and ECM estimates. Baseline model (3.1-3.2) and shocks to real income (4.1-4.2)

Variables	(3.1) LDV FE	(3.2) ECM	(4.1) LDV FE	(4.2) ECM
RHP (L)	0.536***		0.495***	
	(0.0394)		(0.0448)	
RDI	0.410***	0.964***	0.836***	1.088***
	(0.107)	(0.115)	(0.283)	(0.328)
NAIRU (L)	-0.00500	-0.0156***	-0.0188**	-0.0593***
	(0.00453)	(0.00513)	(0.00940)	(0.0105)
Composite real rate (L)	-0.00266**	-0.00207*	-0.00343***	-0.00177
	(0.00106)	(0.00121)	(0.00114)	(0.00132)
Dwelling stock (L)	-0.0167**	0.0101	-0.0244**	-0.0144
	(0.00802)	(0.00865)	(0.0100)	(0.0114)
Pop. density (L)	-0.0652	0.579	2.295**	2.912***
	(0.729)	(0.752)	(0.949)	(0.915)
CPI (L)	-0.00177**	-0.00369***	-0.00172*	-0.00234**
	(0.000806)	(0.000909)	(0.000940)	(0.00110)
Error Correction term (L)		-0.0736***		-0.0635***
		(0.0130)		(0.0243)
Demand shocks and structural factors				
RDI× Banking Supervision			-0.292***	-0.353***
			(0.0892)	(0.104)
RDI × Supply elasticity			0.0214	0.305
			(0.268)	(0.311)
Constant	0.0845**	-0.0282	0.113***	0.0756
	(0.0341)	(0.0369)	(0.0416)	(0.0478)
Total impact (elasticity) of interacted variables on short term price growth at the mean of the structural variables (mean degree of banking supervision= 1.385 over 1975-2006)				
1 % RDI growth			0.43	0.60
Observations	493	507	384	384
Number of countries	20	20	18	18
R-squared (within)	0.440	0.268	0.503	0.326
Still serial correlation? (LM test compared to AR1)	Yes (much less)	Yes (less)	Yes (much less)	Yes (less)

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Everything else equal, the short term elasticity of prices with respect to real disposable income growth is estimated at 0.41 and 0.96 in the LDV and ECM baseline models (3.1-3.2). However, when we allow for different degrees of banking supervision, these results yields diverse effects of growth in disposable income.

CPI inflation and real interest rates have significant and negative effects on yearly RHP growth in all cases and these effects are quite robust to all specifications. Structural unemployment and population density growth is significantly negative when we include the interactions of wage growth with the structural factors and with the expected signs. Supply side factors such as the elasticity of supply identified through changes in income have poor short term explanatory power with respect to RHP growth, consistent with the theory on determinants of short term fluctuations in house prices. The change in the dwelling stock is significant and negative in the baseline models, but this effect disappears when allowing for interactions between income and bank prudence.

The coefficient on the speed of adjustment of yearly RHP growth is about one half in the LDV cases, quite similar to our estimates of the AR(1)-coefficient. The coefficients on the error correction term in the ECM cases which measures yearly speed of adjustment toward equilibrium, are both small in absolute value and highly significant, which we would expect from the persistence detected in price movements. Any given deviation from equilibrium will then take a long time to correct in these specifications.

How does these results compare to other studies? Annett et al (2005) estimate a panel regression on 8 Euro Area countries using fixed effects and ECM and find a short term elasticity of disposable income with respect to real house prices of 0.26 (0.7 in an ECM model), an elasticity of own price of 0.75 and effects of long term interest rates of -0.01 in a FE model. For our estimations, the significant and negative effects of the real interest rate, unemployment and the housing stock supply and the significantly positive effects of income and historic price movements are in line with estimates of longer term effects on price levels found in several studies (see Girouard et al. 2006 p. 11-15 for a survey although most are individual country studies).

To the extent prudential banks capture a credit-effect, though not directly comparable, Almeida et al (2006 p.14 and 17) find positive and significant effects of the maximum Loan-to-Value ratio[13] in a country interacted with income growth on yearly house price changes for 26 OECD countries over 1970-1999. So we may be able to conclude our results in this section conforms with existing findings of the driving factors behind short term price movements in a panel context.

[13] We expect the degree of banking supervision and the maximum LTV ratio to capture opposite effects.

5 The risk of large movements in real house prices

A housing cycle is what happens between the short term movements and long run growth of housing prices

In section 4 a model for the change in house prices from year to year was estimated. Prices were explained by fundamental variables suggested by economic theory. The above quote is not really a quote in the ordinary sense that some influential economist has said it. But there is a tendency in empirical and theoretical work on house price formation to focus on short and long term determinants and then *volatility* or standard deviations of prices for what happens in between. These approaches in turn have in common that they silently assume a *unified* framework for ups and downs. The recent crisis in USA, Ireland and Spain has spurred interest in what causes abrupt falls in house prices. Some speak of bubbles that burst. One idea that has gained attention is that bubbles create asymmetries in the distribution of price changes, abrupt falls being more likely than abrupt increases.

Let us investigate some of the time series properties of a simple linear AR(2) model for the composite (weighted and including 21 countries) OECD real log house price series (p)

$$p_t = \phi_0 + \phi_1 p_{t-1} + \phi_2 p_{t-2} + u_t \quad (5.1)$$

, where $u_t \sim IID(0, \sigma_u^2)$.

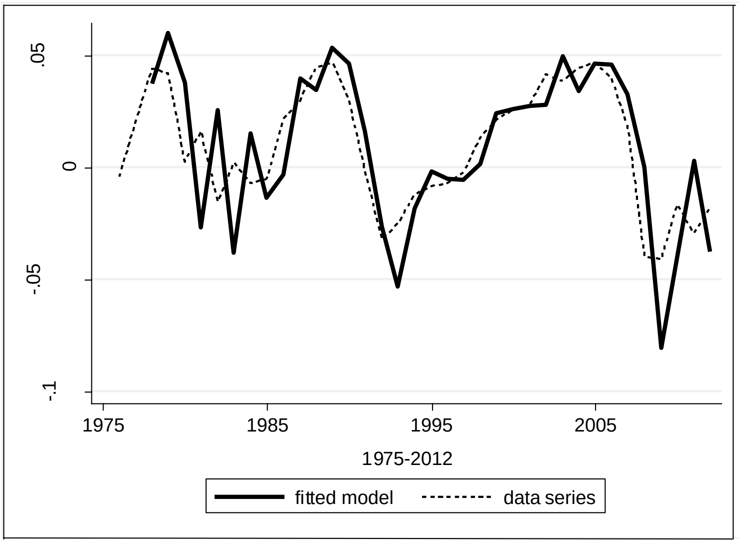
A simple OLS estimation of (5.1) yield the results (5.2) with robust standard errors in parentheses

$$\underset{}{p_t =} \underset{(0.114)}{0.239} + \underset{(0.097)}{1.694 p_{t-1}} - \underset{(0.095)}{0.748 p_{t-2}} \qquad (5.2)$$[14]

[14] The homogenous solution to (5.2) is $p_t = (0.865)^t (A cos(0.979t) + B sin(0.979t))$, for some constants A, B.

Both roots of the corresponding solution for p_t are complex and the long term solution is stable with damped oscillating swings, consistent with the visual impression of a cyclical, mean reverting pattern. Figure 4.1 plots the yearly changes for a fitted model using the parameters in (5.2) and the actual series for the OECD composite real house price yearly growth.

Figure 4.1 Fitted AR(2) model RHP growth



OECD weighted composite. 21 countries

The first thing that should be noted is that the AR(2) model captures the overall movement of the data. However, this fitted linear model has much larger spikes (which would translate into swings in the levels version), a feature that is most evident for negative growth than above the zero-line, especially over the large falls. If rises and falls in prices occur at different speed, the possibility that different economic forces are at work is at least present in the data.

The model from section 3 can in principle be used to estimate the probability of large changes in house prices contingent on the explanatory variables. However, if there are asymmetries or nonlinearities the estimates could be misleading. There is a case for estimating probabilities for large falls and increases in prices more directly, a task we now turn to.

Section 5.1 defines what is going to be meant by large price changes. Section 5.2 defines an econometric model for estimating boom- and bust risks separately. In 5.3 are estimation results from a linear probability model. This allows for an analysis of asymmetry; whether changes in any covariates are driving factors behind booms and busts and if there are signs of any differential driving factors underlying a protracted period of price appreciation and a period of depreciation. Section 5.4 focus on abrupt price falls and estimates a logistic model as an alternative. In 5.5 a forecasting model is estimated and 5.6 fit this model for the US.

5.1 A cut-off value for booms and busts

To qualify as a boom, define a cut-off value of price appreciation of more or equal to 22 % over 3 years (the volatile periods "in between" short and long term growth), which amounts to an average of about 7.33 % yearly growth (H) in real prices. A bust is defined as a 10 % real price depreciation over 3 years as well[15] (L) or 3.33 % yearly decrease. These particular values are chosen since we capture around 15 % of the empirical 3 year price movements in each case, and decide to call these episodes for booms and busts henceforward.

This procedure is chosen for simplicity, more advanced methods such as the Bry and Boschan (1971) cycle-dating procedure used in the literature could be a refinement on this point (described in Harding 2003). The three year criterion may, however, seem more restrictive than it actually is, any boom/bust lasting longer than 3 years will qualify but shorter period cycles will not[16]. Table 5.1 below lists the episodes of real price depreciations satisfying criteria (L)

---

[15] A binary variable for each cut off value is defined where $P_{i,t}^{H} = 1$ if the boom criteria are satisfied and zero otherwise, with a mean value .155. Similarly $P_{i,t}^{L} = 1$ if the bust criteria are met and zero otherwise, with mean value .155, where the 3 years prior to 1978 is excluded since we cannot calculate 3 year changes.

[16] If a country experience a 5 year boom, then this model will evaluate the first three years of the boom, the forth to the second year of the boom etc.

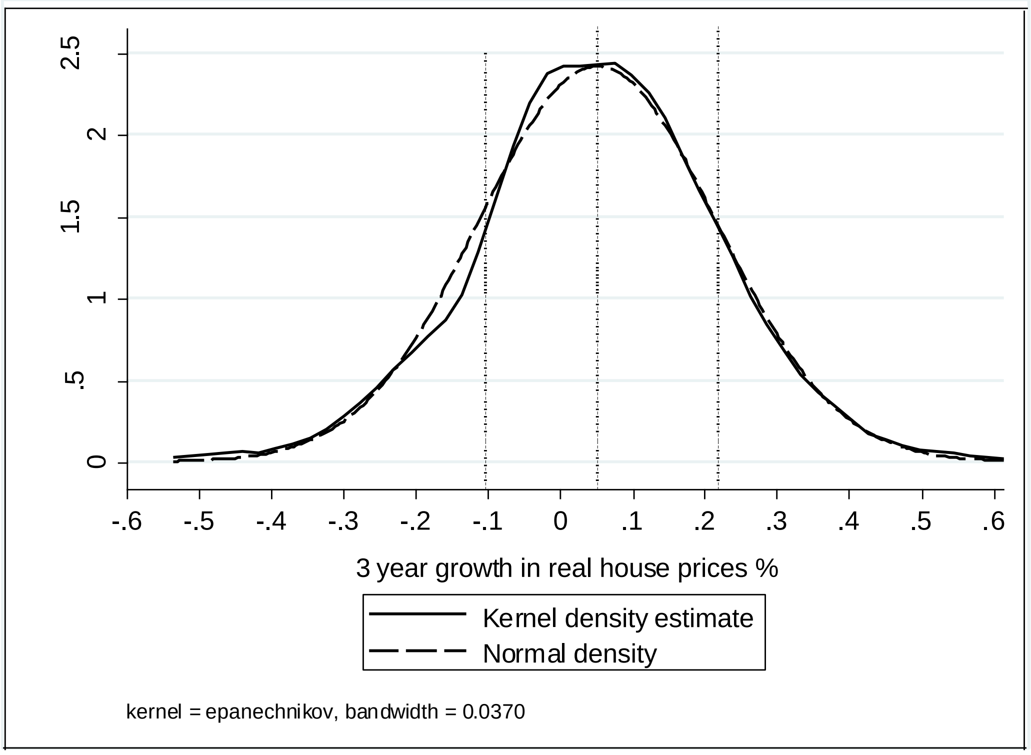
Table 5.1 Episodes of real house price busts by country 1975-2012

The episodes and years in the table satisfy the bust criteria of less than or equal to 10 % real price depreciation the last 3 years.

Country	timespan	total contraction %	episodes	years
1 Australia	-	-	0	0
2 Belgia	’79-‘85	-41.7	1	6
3 Canada	’82-‘85	-33.6	1	3
4 Switzerland	’89-‘98	-45.0	1	9
5 Germany	’82-‘85	-12.0	1	3
6 Denmark	’79-’83,	-26.1		
	’86-’92,	-34.2		
	’07-‘12	-29.0	3	15
7 Spain	’79-’84	-23.5		
	’91-’94	-17.2		
	’07-‘12	-36.9	3	13
8 Finland	’75-’79	-19.7		
	’89-‘95	-56.9	2	10
9 France	’80-‘85	-18.6	1	5
10 UK	’89-’94	-27.6		
	’07-‘11	-16.0	2	9
11 Ireland	’79-’82	-11.0		
	’81-’84	-11.1		
	’06-‘12	-42.0	3	11
12 Japan	’91-’94	-14.0		
	’00-‘06	-24.0	2	9
13 S. Korea	’83-’87	-13.4		
	’90-‘00	-73.8	2	14
14 Netherlands	’78-’84	-58.5		
	’08-‘12	-19.4	2	10
15 Norway	’87-‘93	-40.6	1	6
16 New Zealand	’75-’80	-43.9		
	’07-‘10	-12.0	2	8
17 Sweden	’79-’85	-44.8		
	’90-‘95	-32.5	2	11
18 US	’06-‘12	-28.7	1	6
19 South Africa	’75-’79	-21.0		
	’82-’88	-45.9,		
	’89-‘94	-19.5	3	15
20 Italy	’75-’79	-26.7		
	’81-’85	-20.3		
	’92-’98	-30.0		
	’08-‘12	-18.1	4	18
21 Luxembourg	’79-‘84	-24.9	1	6
Total			**38**	**187**

There were 38 episodes in total where real prices fell more than 10 % over 3 years covering 187 year observations across countries. Denmark, South Korea and Italy have had most 3-year spans with large falls in prices, while Australia is the only country with none such episodes. A similar account of 3 year increases in real prices of more than 22 % (H) lists 47 such episodes. Below we have drawn the Kernel distribution estimate for 3 year change in real house prices along with a fitted normal density

Figure 5.1 The empirical distribution of 3 year price dynamics



The Kernel density for 3-year changes in prices appears normally distributed with more mass for positive values than negative. Average 3-year real price appreciation is 5.2 %. As can be seen in the figure, the booms are more frequent and translate into higher growth rates in our panel than the busts in the extreme cases (there have been episodes with up to 60 % 3-year price appreciations, but only a bit less than 52 % 3-year price depreciations). The cut-off values are marked in the Kernel. In comparison to the AR(2) fitted model for yearly price changes, the overestimation of price falls suggested by figure 4.1 is consistent with the differences in tails in the medium run actual data. Otherwise the distribution is quite symmetrical around the mean.

For our theoretical model, augment equation (5.3) below to apply for three year or lagged three year changes (5.4), which can be interpreted as medium term price appreciations[17]. Then the change in *the growth rate* of new dwellings in $\Delta s_{i,t,t-3}$ is implemented instead of the log new dwellings relative to base in the short term model

$$\Delta p_{i,t} = \beta_0 + \boldsymbol{\beta_1} \Delta \boldsymbol{X}'_{i,t} + \beta_2 \Delta s_{i,t} + \alpha_i + \varepsilon^D_{i,t} \quad (5.3)$$

$$\Delta p_{i,t,t-3} = \gamma_0 + \boldsymbol{\gamma_1} \Delta \boldsymbol{X}'_{i,t,t-3} + \gamma_2 \Delta s_{i,t,t-3} + \zeta_i + \epsilon^D_{i,t} \quad (5.4)$$

5.2 Large price movements –econometric analysis

Redefine the dependent variable as a binary variable equal to 1 if criterion H is met and another equal to 1 if criterion L is met. Then a Linear Probability Model (LPM) version of (5.4) is[18]

$$Prob\left(P^m_{i,t} = 1\right) = \gamma_0 + \gamma_1 \vartriangle logW_{i,t.t-3} + \gamma_2 \Delta NAIRU_{i,t-1,t-4} + \gamma_3 Structural._i * \vartriangle W_{i,t.t-3} + \gamma_3 Structural._i * \vartriangle NAIRU_{i,t.t-3} + \gamma_4 Structural._i * \vartriangle density_{i,t.t-3} + \gamma_5 \Delta S_{t-1,t-4} + \gamma_6 \boldsymbol{Z'}_{t-1,t-4} + \zeta_i + \epsilon_{i,t} \quad (5.5)$$

$$, t = 1, \ldots, T_i, \ \ n = 1, \ldots, N \ , m = H, L$$

,where $P^m_{i,t} = 1$ if $\left(P_{i,t} - P_{i,t-3}\right) \lesseqgtr k$

Equation (5.5) states that the probability of a boom (bust) for a country *i* in year *t* is a linear function of demand side factors whose effect depends on country specific structural factors, the change in the growth rate of the dwelling stock and a vector of controls. The abbreviations have the same interpretations as in section 4. Since we look at past 3 year changes, the covariates are mainly included as past 3 year changes in line with equation (5.4) or changes between 4 years ago and last year to look at the run up year before the boom/bust start-up. The house price growth in the run up year is also included in the vector of controls ***Z.***

[17] The optimal way of including these explanatory variables is not obvious and several methods have been tested.
[18] There are several limitations to the linear probability model discussed in Appendix; however, the model is a natural extension to our baseline model and allows for easy parameter interpretation.

The country specific effect ζ_i can be thought of as the combined effect of omitted country-specific covariates that cause some countries to be more prone to large rises or falls in prices as others, such as home ownership rates or housing investment patterns.

A multiple of specifications have been tested. The three types of demand shocks; to the NAIRU, real disposable income, and demographics interacted with structural factors ($Structural._i$) are included sequentially in the LPM FE/RE framework with "lagged dependent variable[19]". Of course, when the dependent variable is transformed this way, this interpretation is no longer strictly correct.

5.3 Results Boom and Bust risks

Growth to real disposable income is still important in the medium run and indicative of boom and bust probabilities. Other macroeconomic terms have higher power towards boom probabilities than busts. Also, this evidence suggest that supply-side and structural factors matters more in the cases of large price movements than in the general short term models from section 3.

The estimates below indicate some interesting results in direction of symmetry and asymmetry;

a) The baseline models (1.1-2)

In the baseline version corresponding to section 4, whether real house prices rose in the pre potential boom/bust year affects the probability of a boom/bust with the expected signs; such a feature increases the probability of a boom and decreases the probability of a bust. For our other macroeconomic terms, growth to real disposable income in the run-up year and during the potential boom significantly increases the boom- and reduces the bust probability, whereas increased CPI inflation reduces the boom-chances. However, CPI growth is insignificant with respect to the probability of a bust.

An increase in the growth rate of new dwellings in the run-up year and during a potential bust, significantly (at 1 % level) increases the bust-chances *when countries have equal or below*

[19] The model with LDV had least serial correlation in the yearly price dynamics models.

OECD average change in the growth rate of new dwellings[20] and reduces the bust-chances otherwise. This latter feature is consistent with the idea that countries that experience especially rigid supply in the medium term (3 years here) as compared to the OECD average, and the rigidity increases, will have a higher bust-probability and vica versa. Also, according to these results, this effect is asymmetric; the probability of a boom is not affected in either specification below. Andrews (2010 p. 25) find positive but (highly) insignificant effects on overall house price volatility from dwelling investment growth.

b) Structural factors models (2.1-2) Macro-financial linkages

According to these results, negative shocks to demand represented by an increase in the NAIRU have differential effects depending on the level of property tax[21] in the country; An increase in structural unemployment will decrease both the probability of a boom and a bust depending on the level of residential property tax in the country, which suggests a reduction in overall volatility. These results are in line with the findings of Andrews (2010 p. 25) that tax relief on housing (the opposite) increases overall volatility.

c) Structural factors model (3.1-2) Supply-demand linkages and credit effects

If instead we believe structural factors affect booms and busts mainly through income shocks, then the boom-probability following an increase in the growth rate of real income decrease quite sharply if the elasticity of supply is high while such effects are not evident for busts. Instead, and somewhat surprisingly, the bust probability following an increase in the growth rate of income *increases* in the degree of banking supervision.

Daring to propose an explanation to this latter, somewhat counterintuitive result, it *could* result from a) The indicator capturing supervision of banks is positively correlated (the correlation coefficient is 0.5) with credit growth defined as credit issued by deposit money banks and other financial institutions to GDP (WB data). It makes sense that countries increase supervision when inhabitant's debt to income ratios inflates. Or both have increased over time due to expansive financial markets and common OECD standards. We could thus capture credits effects, as intended, but with the reversed sign. b) Higher supervision countries may on average have higher standards towards borrowers. Since wages are rigid downwards

[20] New dwellings are normalized to a base year, so only relative changes matter within each country (See 4 Data). Average 3 year change in the growth rate of new dwellings is -0.053.
[21] It is particularly important to pay attention to possible endogeneity problems related to this property tax measure, which is always included with at least one lag.

we would hardly ever see a fall in income even though prospects are bad. Then cautious banks may hold back on mortgages to the most risky applicants thereby increasing the risk of a bust as credit constrained households are held out of the market.

Andrews (ibid) find that increased banking supervision decreases overall volatility, when evaluated alone, and not combined with a shock to income. These results then conflicts with Andrews' although direct comparison is difficult due to the differences in specifications.

d) Structural factors model (4.1-2) Supply-demand linkages

The impact from demographic shocks measured by an increase in the population density growth rate is somewhat ambiguous, depending on the model being estimated. There is however evidence of decreased overall volatility when increases in population density is combined with high elasticity of housing supply, as this decreases the risk of both a boom and a bust in (4). Andrews find ambiguous results of the supply elasticity on overall volatility, depending on the model and insignificant effects of increases in population density (ibid).

Some general remarks; I am restricting attention toward signs of effects as well as statistical significance due to a couple of nonsensical results regarding magnitudes in this linear model. The most outlandish must be the response in column 3.1 of a 1 % income shock, which supposedly increases the probability of a bust by 3.2. But both signs and significance is robust (there are however some differences in significance levels) to alternative models with more theoretically plausible outcomes bounded on (0, 1). The magnitudes of effects will be discussed in a logistic model in the next section considering risks of price falls only. Remember also the models with structural factors are estimated on a shorter time period (1975-2006) than the baseline models (1975-2012) due to data shortage, and for the same reason cover only 18 countries as opposed to 20. There are thus some problems of comparability.

Results LPM FE models with lagged dependent variable. All variables are in 3 year changes except the structural factors which are in levels and the run up year price growth (RHP L.3). (L) signifies the variable is lagged and panel robust standard errors are in parentheses.

Table 5.2 Estimation results 1 Boom and Bust risks LPM

Dependent variable: Pr(Boom/Bust). Baseline model (1.1-1.2) and shocks to the nairu (2.1-2.2)

VARIABLES	(1.1) LPM Boom	(1.2) LPM Bust	(2.1) LPM Boom	(2.2) LPM Bust
RHP (L.3)	0.888***	-0.386	1.273***	-0.0543
	(0.304)	(0.272)	(0.328)	(0.277)
RDI	1.534***	-0.617*	2.187***	-0.935**
	(0.416)	(0.371)	(0.466)	(0.405)
CPI (L)	-0.0163**	-0.00113	-0.0223***	0.000578
	(0.00739)	(0.00660)	(0.00796)	(0.00663)
Composite real rate (L)	-0.0111*	0.00417	-0.00917	-0.000591
	(0.00653)	(0.00583)	(0.00688)	(0.00572)
NAIRU (L)	-0.0321*	0.0514***	-0.0512	0.0828**
	(0.0187)	(0.0167)	(0.0457)	(0.0400)
Pop. density (L)	0.593	2.887*	0.0638	-1.187
	(1.925)	(1.720)	(1.301)	(1.339)
Dwelling stock (L)	0.0273	-0.233***	-0.0692	-0.217***
	(0.0561)	(0.0501)	(0.0652)	(0.0548)
Demand shocks and structural factors				
NAIRU × Supply Elasticity			0.0758	0.0961*
			(0.0585)	(0.0507)
NAIRU × Property Tax (L)			-0.0401**	-0.0475***
			(0.0171)	(0.0143)
Constant	0.0513	0.0994**	0.0231	0.175***
	(0.0481)	(0.0430)	(0.0392)	(0.0391)
Total impact of interacted variables on boom/bust probabilities at the mean of the structural variables (mean property tax revenues equal to 1.11 % of GDP over 1975-2012, mean supply-elasticity= .703)				
1 % increase in NAIRU			-0.044	0.098
Observations	446	446	404	404
Number of countries	20	20	18	18
R squared	0.143	0.215		
Fixed or random effects	FE	FE	RE	RE
Hausman test P-value (RE efficient under Ho)	0.000	-	0.0002	-

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 5.3 Estimation results 2 Boom and Bust risks LPM

Dependent variable: Pr(Boom/Bust). Shocks to real disposable income (3.1-2) and population density (4.1-2)

Variables	(3.1) LPM Boom	(3.2) LPM Bust	(4.1) LPM Boom	(4.2) LPM Bust
RHP (L.3)	0.963***	-0.283	1.104***	-0.242
	(0.370)	(0.306)	(0.330)	(0.286)
RDI	4.627***	-2.642***	2.276***	-0.934**
	(1.127)	(0.924)	(0.466)	(0.448)
CPI (L)	-0.0118	0.00198	-0.0192**	0.00492
	(0.00849)	(0.00703)	(0.00797)	(0.00671)
Composite real rate (L)	-0.00584	0.00152	-0.00931	-0.000924
	(0.00775)	(0.00642)	(0.00691)	(0.00583)
NAIRU (L)	-0.0442*	0.102***	-0.0433*	0.102***
	(0.0263)	(0.0217)	(0.0230)	(0.0213)
Pop. density (L)	3.943**	-3.215**	3.013*	5.454
	(1.905)	(1.534)	(1.744)	(4.016)
Dwelling stock (L)	-0.0724	-0.0433	-0.0775	-0.206***
	(0.0835)	(0.0690)	(0.0651)	(0.0562)
Demand shocks and structural factors				
RDI × Supply elasticity	-2.028**	0.00407		
	(0.815)	(0.660)		
RDI × Banking supervision	-0.432	0.898***		
	(0.370)	(0.304)		
Pop.density × Supply elasticity			-3.339**	-9.900*
			(1.449)	(5.643)
Constant	-0.0302	0.205***	0.0196	0.203***
	(0.0531)	(0.0431)	(0.0384)	(0.0488)
Total impact of interacted variables on boom/bust probabilities at the mean of the structural variables (mean degree of banking supervision= 1.385 over 1975-2006, mean supply-elasticity= .703)				
1 % RDI growth	3.20	-1.40		
1 % population growth			0.67	-1.50
Observations	318	318	404	404
Number of countries	18	18	18	18
R squared				0.271
Wald Chi_2	71.19	78.82	93.57	
Fixed or random effects	RE	RE	RE	FE
Hausman test P-value (RE efficient under Ho)	-	-	0.000	0.1240

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

5.4 Isolating abrupt price falls

There are some evidence of different influential driving factors involved for our simple predictor of booms and busts. The existence of asymmetries in the housing cycles could make demand relations such as that estimated in section (4) dependent on a sort of switching regime, in which case any prediction problem become a nonlinear one. This would evidently affect empirical results. If a downturn is fundamentally different from expansions[22], a symmetric model would fail to capture both. Asymmetries can arise as i) different magnitude of effects, ii) different driving factors or iii) non-reversed signs of effects, among others.

Based on the estimations reported in section 5.3 one might do better by estimating different models for the boom- and bust risks. If we gather the information about relative significance from the symmetric models, two such general relationships that incorporates this is

$$P\left(P_{i,t}^{\boldsymbol{H}}=1\right)=F(\Delta p_{-t},\Delta w(supplyel.)_t,\Delta cpi_{t-1},\Delta nairu(p.tax)_{t-1},\Delta dens(supplyel.)_{t-1}) \quad (10)$$

$$Pr\left(P_{i,t}^{\boldsymbol{L}}=1\right)=G(\Delta w(banksup.)_t,\Delta nairu(p.tax)_{t-1},\Delta dens(supplyel.)_{t-1},\Delta s_{t-1}) \quad (11)$$

, where t ~ 3 year period, $-t\sim$ year(s) prior to 3 year period, $(t-1)\sim$ run-up year and 2 years into 3 year period and $F(.),\ G(.)\sim$ some functions of the included variables/ functions of variables bounded on $(0,1)$.

The relationships (10) and (11) summarizes the discussion in 5.3. For any country *i* in a year *t* shocks to wages give differential effects on the boom-probability (H) through different supply responsiveness in (10), while for the Bust-probability (L) this effect works through bank prudence. Historic house price movement had much more predictive power towards the risks of a boom than a bust. The nairu combined with property tax enters in both, as does population density growth combined with the elasticity of supply. The general price inflation appears only in (10), while the growth rate of the dwelling stock appears only in (11).

I will focus on the probability of a bust for the remaining, that is, on equation (11). Also, assume there is a relationship between (10) and (11) in the sense that if (10) has been realized (there has been a boom in $-t$), this affects the probability of a bust. Table X below reports

[22] The existence of asymmetries over the business cycle has been widely debated but, to my knowledge, there is no discussion of asymmetries over the housing cycle despite the close connection between the two.

estimation results for a RE Logit model[23] (5) where the odds ratios (OR) are given in the table together with a RE LPM (6) for comparison with earlier estimates, for an estimating model based on equation (11)

Table 5.4 Estimation results Bust risks Logit

Dependent variable: Pr(Bust). RE Logit (5) and RE LPM (6) estimates

Variables	(5) RE_Logit	(6) RE_LPM
	OR	Coeff.
Dummy (=1 if boom before potential bust)	5.106***	.144***
	(2.652)	(0.0385)
RHP (L.3)	0.0924	-0.3027
	(0.337)	(0.276)
RDI	0.968***	-2.728***
	(0.0118)	(0.6757)
NAIRU (L)	6.851***	0.1445***
	(4.734)	(0.2389)
Pop. density (L) (×1000)	0.894*	-0.0043**
	(0.0519)	(0.0018)
Dwelling stock	0.599	-0.0835
	(0.484)	(0.064)
Real disposable income growth × Banking Supervision	1.011*	0.9076***
	(0.00594)	(0.2762)
NAIRU × Property tax	0.548**	-0.052***
	(0.151)	(0.01485)
Pop. density × Supply Elasticity	1.031	0.0005
	(0.0411)	(0.00135)
Constant	0.130***	0.188***
	(0.102)	(0.0385)
Total effect of income shock at banking supervision mean (=1.385)		-1.47
Total effect of shock to structural unemployment at property tax mean (=1.1 % of GDP)		0.102
Total effect of 10 % population density growth		-0.43
Observations	335	335
Number of countries	18	18
LR-test for rho=0: chi_2(1)	1.43	
Wald Chi_2	27.83	132

Standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

According to the estimates, a 3 year real price appreciation of more than or equal to 20 %, increases the probability of a subsequent bust by 14.4 % or the odds of a bust by about 5 compared to a country with no such boom. The estimates are significant at the 1 % level. Otherwise these models summarize the findings from our previous section. The results are

[23] Both the Logit and an equivalent Probit model estimation reports a log likelihood of -141.

quite robust to different specifications, as the comparison of the LPM and Logit estimates broadly confirms in columns (5) and (6).

The model (11) explains the risk of large falls in real house index prices over a period of 3 years in 18 OECD countries as a function of changes in structural and demand side factors and historic price behavior. The particular specification of dynamics, however, limits inference of driving factors to a rather contemporaneous one. The probability of a bust is explained by random country specific effects and changes in the confounders *during* and only including the run up year of any potential bust. This perspective is valuable, but more so if the model has *forecasting* abilities.

5.5 Forecasting

From the observed rigidities and cyclicality of house prices and when prices are thought to be misaligned with *fundamentals* (such as income or unemployment) one often talks of build-ups in the housing market. The development in the overall economy and the history of prices over a period of time, such as 3 years or more may then contribute to such build-ups. In this process we will expect the risk of large falls in prices to gradually increase. Here we consider the (in sample) forecasting abilities of model (11) to predict the probability of falling prices next year based on a 4 year history of development in the confounders and evaluate the model against data. Will the model be able to predict e.g. the protracted fall in prices starting in the US in 2006?

The dichotomous dependent variable is set to one if real prices fall more than or equal to 3.3 percent the coming year[24], the equivalent of the 3 year criterion in the previous sections. I use the logistic model due to its theoretical advantages, and define a model (12)

$$E\left(Pr\left(P_{i,t+1}^{F}=1\right)\middle|\boldsymbol{x},\alpha_i\right)=\frac{e^{\boldsymbol{x\beta}}}{(1+e^{\boldsymbol{x\beta}})}$$

$$,\boldsymbol{x\beta}\coloneqq\beta_0+\beta_1 DBoom+\beta_2\Delta P_{i,t-1,t-4}+\beta_3\vartriangle logW_{i,t.t-3}+\beta_4\Delta NAIRU_{i,t-1,t-4}+\beta_5\Delta Density_{i,t-1,t-4}+\beta_6 Banksup._i*\vartriangle W_{i,t.t-3}+\beta_7 Prop.tax_i*\vartriangle NAIRU_{i,t.t-3}+\beta_8 SupplyEl._i*\vartriangle density_{i,t.t-3}+\beta_9\Delta S_{t-1,t-4}+\alpha_i+\epsilon_{i,t}\quad(12)$$

, *F*= Falling prices next period

---

[24] This criterion is met in 19 % of the country year observations

The interpretation of the variables is the same as before and *E* is the mathematical expectations operator. The right hand side of this first expression is the CDF of a logistic distribution and all coefficients should be interpreted as conditional on the country specific effect.

Below are estimation results reported by the odds ratios together with the average marginal effects (AME) evaluated at $\alpha_i = 0$. The coefficients on historical price movements are both large in magnitude and highly significant, resulting in large increases in the odds of price falls following a 3 year increase in prices. A test of the joint insignificance of all other 7 coefficients (from income growth onwards in table 5.5) rejects at 0.05 % ($\chi^2(7) = 25.97$). We are thus not only capturing own price effects in (12)

Table 5.5 Estimation results Forecasting price falls

Dependent variable: Pr(falling prices). RE Logit odds ratios and AMEs

Variables	(7) RE Logit	AME
	OR	
Dummy (=1 if boom before potential bust)	3.022***	0.1394***
	(1.041)	(.0410)
D3. RHP growth %	72.26***	0.5396***
	(98.56)	(1589)
D3. Real disposable income growth %	1.000	-0.000
	(0.00665)	(.0008)
D3. Change in NAIRU % (L)	2.372***	0.109***
	(0.669)	(.0311)
D3. Population density growth % (L) (×1000)	0.937**	-0.008**
	(0.0263)	(0.0034)
D3. Change in dwelling stock growth rate	0.827	-.0239
	(0.494)	(0.075)
Real disposable income growth × Banking Supervision	0.995*	-0.00065*
	(0.00326)	(0.0004
NAIRU × Property tax	0.7991*	-0.0283*
	(0.115)	(0.017)
Population density growth × Supply Elasticity	1.033*	0.0041*
	(0.0203)	(0.002)
Constant	0.282***	
	(0.129)	
Observations	335	
Number of countries	18	
LR-test for rho=0: chi_2(1)	1.56	
LR Chi_2	53.06	
Rho (intra-class correlation)	0.083	

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

How well does this model predict historical episodes? The significant parameters from the RE Logit estimation are used to calculate probabilities of falling prices at the eve of a bust satisfying criteria L and compared to country specific averages evaluated at $\alpha_i = 0$ in table 5.6. Compare the tipping points with the episodes listed in table 5.1.

Table 5.6 Bust probabilities at the eve of protracted price falls 1975-2006

The episodes satisfy the bust criteria of more than or equal to 10 % real price depreciation the next 3 years.

Country	**Tipping point(s)** ‘year	**Pr(Bust) (**mean**)**	**SD from mean**	**SD**
1 Australia	none			
2 Belgia	‘79	**.809** (.256)	2.99	.185
3 Canada	‘81	**.486** (.136)	3.22	.109
4 Switzerland	‘89	**.159** (.150)	0.07	.137
5 Germany	-			
6 Denmark	’79 and	**.672** (.268)	2.10	.192
	‘86	**.580** (.268)	1.63	
7 Spain	‘82[25] and	**.801** (.339)	1.70	.273
	‘91	**.658** (.339)	1.17	
8 Finland	‘88	**.788** (.386)	1.63	.247
9 France	‘79	**.580** (.175)	2.55	.159
10 UK	’89 and	**.250** (.219)	.18	.171
	‘05	**.560** (.219)	1.99	
11 Ireland	‘05[26]	**.044** (.025)	1.19	.016
12 Japan	‘91	**.330** (.209)	1.51	.080
13 S. Korea	-			
14 Netherlands	’79 and	**.839** (.193)	3.08	.208
	‘05	**.500** (.193)	1.48	
15 Norway	‘87	**.520** (.170)	2.33	.150
16 New Zealand	-			
17 Sweden	’79 and	**.542** (.299)	1.60	.152
	‘90	**.546** (.299)	1.625	
18 US	’79 and	**.554** (.191)	2.52	.144
	‘05	**.256** (.191)	1.17	
19 South Africa	-			
20 Italy	’81 and	**.484** (.330)	.92	.167
	‘92	**.495** (.330)	.99	
21 Luxembourg	-			

The model (12) was estimated over the histories of the countries in question so we would expect to get higher probabilities on average before an adverse price movement. But it is one

[25] First available data for Spain,, while the tipping point is around 1979. This means we here evaluate the probability in the midst of a bust.

[26] Due to data shortage the episode starting in ’79 is not included in the table. This shortage also makes the mean and SD probability of high-volatility Ireland questionable since the calculations are based on a short, perhaps unrepresentative sample.

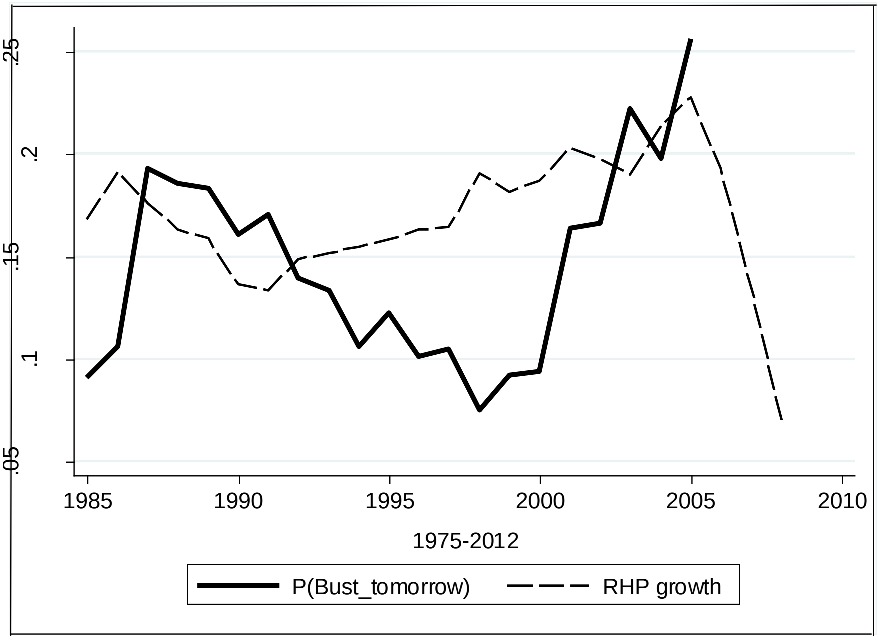
thing to say the probabilities are higher on average and quite another to expect it to hold at a single point in time for a country. In addition, the dependent variable in (12) is not taking into account the cut-off measure (<=-10 % over 3 years) in real prices but nonetheless serves as a proxy in that the aggregated series show clear state dependence between subsequent years (ref. section 2).

The overall picture is that the model predicts many of the significant downturns experienced by a number of countries over the last decades. For Belgium, Canada, France, the Netherlands ('79 case) and the US ('79 case) the calculated probabilities are between 2.5-3.2 standard deviations away from their respective means. The model does not fare well in predicting the downturns in Switzerland ('89) and the UK ('89) however, where such probabilities would be regarded as 'business as usual' on any interesting scale.

5.6 The case of US

The crash in the US housing market in 2006 had widespread consequences for the world economy and has since directed increased attention among economists and policy makers towards the importance of the functioning of the housing market in connection to the real economy. Figure 5.2 display the evolution of our fitted forecasting model for the US along with the actual house price growth

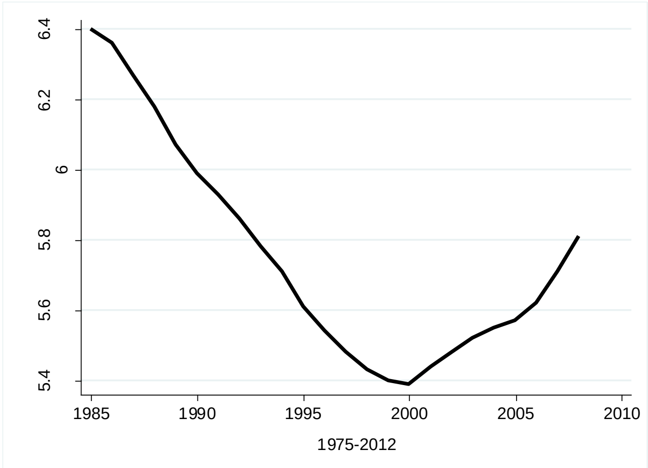
Figure 5.2 The evolution of the forecasted risk of a Bust for the US



Predicted probabilities based on the results in table 5.5. A constant is added to the actual real house price growth series for visual comparison.

As can be seen in the figure, the bust risk increased quite sharply between 2000 and 2005, interrupted only by a slight fall in 2004. In 2005 it reaches a high of .256, which is 1.17 standard deviations away from the mean (table 5.6) based on the longer time period 1975-2006. This upward movement of risk is driven by a long period of increasing house prices, increases in structural unemployment (figure 5.3), a decline in the population density growth rate and falls in the growth rate of disposable income for some, but not all years (that is, the 3 year development up until the particular year).

Figure 5.3 The US NAIRU over 1985-2008



Data: OECD

Figure 5.3 show that structural unemployment increased quite markedly in the US between 2000 and 2008, and contributed to a noticeable increase in the risk in the housing market in my model.

6 Conclusion

The purpose of this thesis has been to investigate empirically *how* differences in structural and policy factors affects the house price responsiveness of shocks to demand in the short run and in the cases of abrupt movements in house prices.

To this end several, relatively parsimonious, models founded in economic theory are presented, each aiming at addressing different aspects of some of the questions involved. Each model builds on the former, from start to end.

The advanced country perspective is chosen since the national or aggregate level is the most natural unit for comparison of differences in regimes, and at the same time similar enough to allow any comparison at all. It should be noted, however, that aggregate information at the country level might hide important regional differences within countries.

A parallel purpose has been to explore a non-unified treatment of booming and busting house price periods and provide empirical evidence of asymmetries and similarities, which is an undeveloped area of economic research.

In a short term (yearly) model I find evidence of diverse effects of shocks to households' income depending on the standards towards monitoring of banks and financial institutions.

Evaluating periods of sharp price falls and increases separately, these results indicate that the responsiveness of housing supply coupled with rises in households' income increases the risk of a boom while for the risk of a bust this effect works through bank monitoring. Adverse demand shocks such as increases in structural unemployment, reduces overall volatility and more so the higher is the residential property tax.

A separate model for price falls is tested against data through a forecasting exercise. The models' ability to predict a bust is improved when we allow for interactions between determinants of demand and structural factors. At the eve of a bust, this model signal between .92 and 3.22 standard deviations higher risks in the housing market than the average or mean risk for 20 out of 22 episodes, and fails in 2 cases.

Still some questions remain. This analysis has been almost silent as to *why*, from the perspective of theory, there might exist nonlinearities over the housing cycle. It is also a debate whether house prices are predictable in the first place. The determinants of housing prices is a complex universe and one must deal with intractable interactions between shifting expectations, a battery of policy interventions, scarcity of land, the dual role of housing as an asset and a necessity and inventive mortgage markets, to name some. All models and estimations presented here are subject to such caveats.

Going forward I would propose to develop richer models and employ more refined methods of dating procedures of housing cycles and advocate that efforts are put into high quality data for cross country determinants of the dynamics of housing prices.

Bibliography

Abiad, A., E. Detragiache and T. Tressel (2008), “A New Database of Financial Reforms”, IMF, *Working Paper/08/266.*

Almeida, H., L. Crocker and M. Campello (2005), “The Financial Accelerator: Evidence from International Housing Markets”, AFA 2007 *Chicago Meetings Paper*.

AEI (April 2012), “Free Fall: How Government Policies Brought Down the Housing Market”, *Financial Services Outlook*

Andrews, D. (2010), “Real House Prices in OECD Countries: The Role of Demand Shocks and Structural and Policy Factors”, OECD *Economic Department Working Papers,* No. 831.

Annett, A. et al. (2005), “Euro Area Policies: Selected Issues”, IMF *Country Reports* No. 05/266.

Beck N. (2006), “Time-Series-Cross-Section Methods”, Draft of June 5, 2006.

Beck N. and J.N. Katz (2011), “Modelling Dynamics in Time-Series-Cross-Section Political Economy Data”, *Annual Review of Political Science* 2011.

Best, R. (2008), «An Introduction to Error Correction Models», Oxford Spring School for Qualitative Methods in Social Research.

Building Society Association (BSA 2007), “House Price Expectations: An insight into how people think about property purchase”.

Canarella, G., S. Miller and S. Pollard (2010),”Unit roots and structural change: An application to US House Price Indices”, *Working Paper* 2010-04R, University of Connecticut.

Catte, P., N. Girouard, R. Price and C. André (2004), “The contribution of housing markets to cyclical resilience”, OECD *Economic Studies* No. 38, 2004/1.

DiPasquale D. and W. Wheaton (1992), “The Markets for Real Estate Assets and Space: A Conceptual Framework”, *Journal of the American Real Estate and Urban Economics Association,* V20, 1: p.181-197.

Englund P. and Y. Ioannides (1992), “The Dynamics of Housing Prices: An international perspective”, *Economics in a Changing World,* Volume 3, ch. 10, p.175-197, International Economic Association.

Englund P. and Y. Ioannides (1997), “House Price Dynamics: An international empirical perspective”, ipricjhe.tex.

Greene, W.H, “Econometric Analysis”, 7. Edition, *Pearson Education Limited.*

Igan D. and P. Loungani (2012), “Global Housing Cycles”, *IMF Working Paper* WP/12/217.

Mack, A. and E. Martinez-Garzia (2011), "A Cross-Country Quarterly Database of Real House Prices: A Methodological Note", Federal Reserve Bank of Dallas, Globalization and Monetary Policy Institute, *Working Paper No. 99.*

Meen, G. (2002), "Why do Mortgage Markets Matter", *Economic Outlook,* Volume 24, Issue 4.

Morley J. et al. (2009), "Reproducing Business Cycle Features: How important Is Nonlinearity Versus Multivariate Information?", *Wesley Economic Working Papers* No: 2009-003.

OECD (2006), "Has the rise in debt made households more vulnerable?", *Economic Outlook* No.80.

OECD (2011), "Housing and the Economy: Policies for Renovation", *Economic Policy Reforms 2011 Going for Growth.*

Poterba J. M. (1991), "House Price Dynamics: The Role of Tax Policy and Demography", *Brookings Papers on Economic Activity,* 2:1991.

Poterba J. M. (1992), "Taxation and Housing: Old questions, new answers", *NBER Working Papers Series* No. 3963.

Romer D., "Advanced Macroeconomics", 4. Edition, MHID.

Rødseth A. (1987), "Bustadsmarknaden –utviklingstrekk og verkemåte", *Sosialøkonomen,* nr.11 1987.

Sánchez A. and Å. Johansson (2011), "The price responsiveness of housing supply in OECD countries", OECD *Economic Department working papers* No.837.

Wooldridge, J.M. (2010), "Econometric Analysis of Cross Section and Panel Data", 2. edition, MIT Press.

Wyman M. (2003), "Study on the Financial Integration of European Mortgage Markets", European Mortgage Federation.

Appendix

Assumptions underlying the econometric modeling and tests

The baseline model and extensions

The first choice to make is whether to estimate the joint model consisting of the demand and supply or investment system. This is especially relevant if there is a close relationship between the two, and simultaneous equation methods will be appropriate. However, from the observed slow or insufficient degree of housing investments in many countries with respect to price changes as of making the housing market clear at current prices, since our scope is modelling short to medium term dynamics, it is theoretically plausible and fairly common practice to assume the demand side is most likely to drive prices *given* supply side factors such as the elasticity of supply. So we choose the inverted demand regression model aiming at the proper conditioning and functional form.

The *temporal dominance* calls for methods that say more about the time series structure of the data. But since the panel aspect is likely also to matter, at a minimum, if choosing constant parameter values and joint estimation as opposed to finding parameter values for each country or allowing for heterogeneity more explicitly, we want to allow for some unobserved individual effects. The baseline unobserved effects regression model can be written generally, for a specific country-year-observation, as

$$y_{i,t} = \boldsymbol{X'}_{i,t}\boldsymbol{\beta} + \alpha_i + \varepsilon_{i,t} \quad (A.1)$$

,where $y_{i,t}$ is a scalar, the observed regressors $\boldsymbol{X'}_{i,t}$ is $(K+1) \times 1$,the unobserved α_i is a scalar and $\varepsilon_{i,t} = \rho\varepsilon_{i,t-1} + \epsilon_{i,t}$ as defined in section 4.

According to Beck (2005), when data are annual, short lags of the AR (p)-process are often sufficient, while e.g. quarterly data would call for examination of higher order processes.

Next we need to examine the nature of the unobserved effects. Fixed Effects (FE) and Random effects (RE) estimation rely on strict exogeneity of the idiosyncratic error conditional on the unobserved effect and all T regressor-vectors for consistency $E\left(\varepsilon_{i,t}\middle|\alpha_i, \mathrm{X}_i\right) = 0 \quad, t = 1, \dots, T$, where $\mathrm{X}_i = \boldsymbol{X}_{i,1}, \dots, \boldsymbol{X}_{i,T}$. For the within estimator used in the FE-case, because of time demeaning we lose N=21 degrees of freedom. Both models are estimated using Statas "Xtregar"-function, which in short involves first transforming the data to remove the AR(1)-component and then use the within-transformation to lose the fixed effect in the FE-case or use the Baltagi-Wu GLS estimator for the RE-case. Both estimation methods allow for unbalanced panels.

Also note that A.1 is in fact a dynamic model (ADL with unobserved component)

$$y_{i,t} = \boldsymbol{X'}_{i,t}\boldsymbol{\beta} + \alpha_i + \rho\varepsilon_{i,t-1} + \epsilon_{i,t} = \boldsymbol{X'}_{i,t}\boldsymbol{\beta} + \rho y_{i,t-1} - \boldsymbol{X'}_{i,t-1}\boldsymbol{\beta}\rho + \upsilon_i + \epsilon_{i,t} \quad (A.1)'$$

Hence we could instead use a lagged dependent variable (LDV) model with IID errors and thus make the dynamics part of the model. ; in the AR(1)-case we obtain different dynamics (impulse responses (IRF)) from the observed and unobserved parts of the model. The partial effect on prices from one of the observables will look like a spike, as in the static case, while the unobserved factors have a declining geometric IRF at rate ρ. The LDV has an IRF for both that has a declining geometric form. The distinction with the AR (1)-case is then that we assume the effects of all variables have impacts that die out exponentially, also the unobserved parts (Beck 2006).

Subtract $y_{i,t-1}$ from both sides of $(A.1)'$ and add and subtract $\boldsymbol{X'}_{i,t-1}\boldsymbol{\beta}$ from the RHS and get

$$\Delta y_{i,t} = \Delta\boldsymbol{X'}_{i,t}\boldsymbol{\beta} + (\rho - 1)\left(y_{i,t-1} - \boldsymbol{X'}_{i,t-1}\boldsymbol{\kappa}\right) + \upsilon_i + \epsilon_{i,t} \quad (A.2)$$

The error correction model is popular in studying house price dynamics (E.g. Sanchez and Johansson 2011, Catte et al. 2004). As we have seen this model offers an interpretation of the long term equilibrium notion of house prices and short term fluctuations. However, in ordinary ECM estimations one loses the country specific heterogeneity. If we want to derive panel estimates, we need to make assumptions and suitable handle the υ_i as discussed above in the FE or RE cases. Below are the long term levels estimates used to calculate the error correction term in (4.2)

Table A1 Long term estimates ECM model (4.2)

Variables	**FE_model**
RDI	1.659***
	(0.0973)
RDI ×Banksup.	-0.00636**
	(0.00285)
RDI ×SupplyEl.	-0.783***
	(0.0732)
Pop. density	1.165***
	(0.157)
Nairu	-0.0213***
	(0.00662)
CPI	-0.00439
	(0.00387)
C. real rate	-0.00292
	(0.00352)
New dwellings	0.177***
	(0.0251)
Constant	-0.0348***
	(0.00777)
Observations	396
R-squared	0.769

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

The LDV version of A.1 is

$$y_{i,t} = \theta y_{i,t-1} + \boldsymbol{X}'_{i,t}\boldsymbol{\beta} + \alpha_i + \varepsilon_{i,t} \quad (A.3)$$

Estimating $(A.3)$ by fixed effects we must pay attention to potentially biased estimators. Under the WE-transformation (denoted by ¨), the lagged dependent variable become $\ddot{y}_{i,t-1} = y_{i,t-1} - \frac{1}{(T-1)}(y_{i2} + \cdots + y_{iT})$, while the errors become $\ddot{\varepsilon}_{it} = \varepsilon_{it} - \frac{1}{(T-1)}(\varepsilon_{i2} + \cdots + \varepsilon_{iT})$. Since the $y_{i,t-1}$ term in $\ddot{y}_{i,t-1}$ correlates negatively with the $\frac{-1}{(T-1)}\varepsilon_{i,t-1}$ in $\ddot{\varepsilon}_{it}$ while, the $\frac{-1}{(T-1)}y_{it}$ and ε_{it} terms also moves together, the regressor and error are still correlated after transformation. But since T is quite large (37), the $\frac{-1}{(T-1)}\varepsilon_{i,t-1}$ and $\frac{-1}{(T-1)}y_{it}$ terms tends to zero. Thus when using a lagged dependent variable, the FE-framework have some satisfactory asymptotic properties. Beck and Katz demonstrate this by running Monte Carlo experiments to compare OLS estimation of a TSCS model with lagged dependent variable and fixed effects to the Anderson-Hsiao estimator (IV-approach) with T > 20 and find that OLS outperform Anderson-Hsiao (Beck & Katz 2009).

The linear probability and logistic model

Redefine the dependent variable in (1) as a binary variable. If we assume strict exogeneity of the errors, the conditional probability that y equals 1 is $\Pr(y_{i,t} = 1 | \boldsymbol{X'}_{i,t}, \alpha_i) = E(y_{i,t} | \boldsymbol{X'}_{i,t}, \alpha_i)$, that is the regression function. The most obvious problem of assuming linearity in a probabilistic measure is the possibility of obtaining values outside $[0,1]$ as we see in section 5. Also, standard errors are generally heteroscedastic, and this should be corrected for by using standard errors robust to such heteroscedasticity (Cameron, Trivedi (2009)). It is however a straight forward approach to estimating marginal effects and allows for controlling for fixed effects. The saneness of the LPM ultimately boils down to whether or not the data generating process share the same properties, hence if boom-and bust-probabilities are linear in the parameters.

The RE Logit model specifies that $\alpha_i \sim N(0, \sigma_\alpha^2)$. and a standard logistic distribution for the individual error term. The margins command in Stata computes marginal effects at $\alpha_i = 0$ for identification , and it should be checked whether this really is a representative evaluation point. Alternatively the margins must be interpreted with caution. The estimation is conducted using 20 integration points (the Stata default is 12). Since the response in this analysis is scewed to the left (more zeroes than ones), a complementary log-log model is also tested for. It turned out the log likelihood is -141 as well (-141 for both the probit and the logit model) and produced almost identical fitted probabilities for the case of US, so there is no trade-off here.

Test of the integrated properties of log real house prices

The model that is tested in the CMR AO-case is

$$p_t = \mu + \delta_1 DU_{1t} + \delta_2 DU_{2t} + \sum_{i=1}^{6} \omega_{1i} DT_{b1,t-i} + \sum_{i=1}^{6} \omega_{2i} DT_{b2,t-i} + \rho \hat{p}_{t-1} + \sum_{i=1}^{6} \theta_i \Delta \hat{p}_{t-i} + c_t$$

,where p_t is log of real house prices in period t, $DU_{mt} = 1 \; if \;\; T > T_{bm}$ and 0 otherwise, $m = 1,2$ captures the structural breaks and ρ is the unit root test-coefficient. Including lags of $\widehat{\Delta p}_{t-i}$ allow for serial correlation up to a lag of 6 years[27].

[27] The value of 6 years is chosen by looking at average length of cycles and then adding about a year.

Table A2 Test for unit roots with structural breaks', AO-model. Critical value for $(\rho - 1)$ t-stat is -5.49

country	Optimal breakpoints	$(\boldsymbol{\rho - 1})$	t	Break 1	t	p-value	Break 2	t	p-value
AUS	1990, 2004	-0.634	-4.74	0.398*	6.84	0.000	0.555*	7.83	0.000
BLE	1991, 2002	-0.372	-3.09	0.303*	5.94	0.000	0.445*	7.72	0.000
CAN	1990, 2003	-0.530	-3.18	0.095*	2.43	0.002	0.475*	10.46	0.000
CHE	1987, 1994	-0.652	-3.17	0.24*	5.55	0.000	-0.24*	-5.10	0.000
DEU	1986, 2004	-0.564	-4.96	-0.09*	-6.14	0.000	-0.14*	-7.92	0.000
DEN	1999, 2007	-0.515	-4.11	0.479*	8.16	0.000	0.086	1.05	0.302
ESP	1992, 2006	-0.386	-5.15	0.514*	6.01	0.000	0.387*	3.31	0.002
FIN	1987, 2001	-0.317	-3.70	0.191*	3.26	0.003	0.410*	6.70	0.000
FRA	1992, 2007	-0.382	-4.00	0.297*	4.64	0.000	0.461*	4.87	0.000
UK	1985, 2001	-0.825	-4.79	0.483*	9.64	0.000	0.696*	13.89	0.000
IRL	1993, 2000	-0.678	-4.54	0.474*	6.09	0.000	0.709*	8.47	0.000
JAP	1988, 2001	-0.251	-2.19	0.260*	6.25	0.000	-0.34*	-7.73	0.000
KOR	1995, 2004	-0.341	-3.27	-0.37*	-6.87	0.000	0.106	1.59	0.120
NED	1993, 2000	-0.457	-1.37	0.334*	4.48	0.000	0.455*	5.67	0.000
NOR	1987, 2001	-0.435	-3.58	0.113	1.89	0.068	0.524*	8.34	0.000
NZ	1998, 2006	-0.425	-4.51	0.467*	6.95	0.000	0.332*	3.74	0.001
SWE	1983, 2002	-0.197	-3.44	-0.20*	-3.44	0.002	0.562*	10.17	0.000
USA	1979, 2007	-0.134	-1.75	0.182*	2.46	0.025	0.182*	2.45	0.026
ZAF	1987, 2005	-0.689	-4.97	-0.35*	-5.55	0.000	0.658*	8.50	0.000
ITA	1990, 2004	-0.838	-4.64	0.195*	4.79	0.000	0.224*	4.56	0.000
LUX	1991, 2003	-0.530	-4.33	0.522*	8.75	0.000	0.479*	6.86	0.000
Composite	1992, 2007	-0.290	-4.17	0.158*	4.86	0.000	0.106*	2.21	0.034

' the model used is Clemente-Montañés-Reyer (1998) unit root test with structural breaks. The null is that the series has a unit root with structural breaks against the alternative of stationarity with breaks. Structural breaks are treated as endogenous in the model. The AO-model captures a sudden change in the mean of the series from inspection of the graphs. The alternative within this framework would be the IO-model with gradual shifts.
* indicates the structural breaks suggested by the test is significant at the 5 % level (which is all countries less 3).

Real disposable income

The Im-Peseran-Shin test tests the null that all countries log incomes have a unit root where the autoregressive parameter ρ_i is allowed to vary between countries against the alternative that *some* but not all have unit roots, and seem appropriate for our analysis.

Table A3 Test statistic

```
Im-Pesaran-Shin unit-root test for lrdinc
-----------------------------------------
Ho: All panels contain unit roots          Number of panels  =     22
Ha: Some panels are stationary             Number of periods =     38

AR parameter: Panel-specific               Asymptotics: T,N -> Infinity
Panel means:  Included                                    sequentially
Time trend:   Not included                 Cross-sectional means removed

ADF regressions: 1.91 lags average (chosen by AIC)
------------------------------------------------------------------------
                    Statistic      p-value
------------------------------------------------------------------------
 W-t-bar             -0.3789        0.3524
------------------------------------------------------------------------
```

Including lags allow for serial correlation, the number of lags is determined by the AIC-method. We use demeaning to mitigate the effect of correlation in income between countries. The null of a unit root in all countries is not rejected at the 1 % level using this specification (t-stat is approximately standard normal under the null). The test require T grow larger than N sequentially.